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Regional mobility in the European Union*

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Abstract

This article examines mobility in the regional distribution of per capita income in the European Union between 1977 and 1999. The methodology used to investigate this issue combines a series of measures taken from the literature devoted to the dynamic study of personal income distribution with a non-parametric analysis. The results obtained show limited mobility in the distribution considered, and a decline in mobility over time. The empirical evidence presented indicates, moreover, that mobility patterns vary as a function of the regional development level. The analysis carried out also highlights the important role played in explaining changes in the regional relative positions by variables such as the initial per capita income, the share in total employment of agriculture, advanced services and non-market services.

Key words: Mobility, per capita income, regions, European Union.

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1 Introduction

In recent years, the issue of territorial imbalances in the European Union (EU) has been examined in numerous studies from a variety of different approaches[1]. There are various reasons for the amount of interest surrounding this question. Among them is the fact that economic growth theory has advanced greatly over the last two decades, coinciding with the introduction of endogenous growth models in the mid eighties. Another, the need to reduce disparities in terms of development levels across the various European regions, is directly related to some of the basic principles behind the forming of the Union, especially since the introduction of the Single Act and the Maastricht agreements. In particular, one of the specific assumptions of the European integration programme is that it will drive the growth of all Member States, thereby increasing economic and social cohesion[2].

Most of the articles dealing with the analysis of regional per capita income disparities in Europe apply the concepts of *sigma convergence* and *beta convergence*, introduced by Barro and Sala-i-Martin (1991, 1992), combining the information provided by various dispersion statistics with the estimation of convergence equations. However, as Quah (1993, 1996a, 1997) has repeatedly pointed out, not only does this approach raise a number of econometric problems, it also fails to capture a series of potentially interesting issues relating to the dynamics of the distribution in question. In particular, this type of analysis does not consider the possibility of regions modifying their relative positions over time, and thereby neglects the whole issue of intradistributional mobility.

As an illustration of the relevance of questions relating to the analysis of distribution dynamics, let us consider the following example. Let us assume that we have information for a period of several years on regional incomes and populations in two given countries, A and B, each of which is in turn divided into two regions with exactly the same size of population. To eliminate from the analysis the effects of population shifts, let us also suppose that there is no change over time in the distribution of the population share in either of the two countries considered. In both A and B, the per capita income of one of the two regions is exactly twice that of the other region, and this situation remains unaltered for the whole of the period considered. There is, however, one major difference between these two countries. A is characterized by a high degree of regional mobility, such that, every year, its two regions switch positions. The situation in B, however, is different in that the relative positions of its regions remain constant year on year. The type of analysis commonly found in the literature is essentially static in its approach, since it is based on cross sectional information, so that it will reveal no appreciable difference between A and B. In fact, given that there is no change over time in the cross sectional structure of the per capita income distribution of either of the countries, any inequality index that satisfies the properties of symmetry and scale independence will give exactly the same value for A and B throughout the period considered[3].

This example highlights the need to supplement standard inequality studies with additional data relating to the mobility of the distribution under analysis. It is precisely this issue that the present article aims to address. Our objective is to analyze mobility in the regional distribution of per capita income in the EU from 1977-1999. By this we hope to contribute to the knowledge of the nature of observed territorial imbalances in the European context, with a view to drawing some type of conclusion that might be of use to regional policy makers within the Community. For indeed, if a given level of inequality were found to be associated with a low degree of mobility, this might indicates that regions are becoming set in their relative positions. If so, this would reinforce the need for an active policy to reduce regional disparities. If, however, the results of the analysis suggest that existing inequality can be largely explained by the variability of regional incomes, regional policy makers would need to take steps to offset adverse economic cycle effects, and let traditional convergence policies take second place.

One of the main innovations in this study relates to the instruments it uses to analyze regional mobility. In contrast to the few articles that have so far dealt with this issue in the European context[4], our working method is fundamentally based on the calculation of a set of measures of the kind used in the dynamic study of personal income distribution. However, since our unit of reference is the region and not the individual, we will proceed by introducing population as a further dimension of the analysis. Thus, the indicators resulting from our calculations will be statistics weighted by the population share of each region, though, in theory, we could take into consideration any variable that were representative of the economic size of the various geographical areas under analysis (income share, surface area, etc.)[5]. Surprisingly, this is an approach that has so far received very little attention in the literature devoted to the analysis of territorial imbalances. This is no doubt due, in part, to the obvious limitations of the theoretical and empirical basis for the analysis of intradistributional mobility[6]. In any event, in order to test the robustness of our results, we will perform a parallel study of mobility in the regional distribution of per capita income using the non-parametric methodology presented by Quah (1996a, 1997). Finally, we will examine the explanatory elements of detected patterns by means of different regression models.

For an analysis of the kind we wish to conduct, it is necessary, furthermore, to obtain a representative sample of the various economies within the context under study while also covering a long enough time period. We have accomplished this by using the

Cambridge Econometrics regional database which has enabled us to employ statistical data on 197 NUTS2 regions for the period between 1977 and 1999[7].

This article is structured into six sections as follows. Sections 2 and 3, which follow this introduction, examine the level and evolution of mobility in the regional distribution of per capita income in the EU using several complementary approaches. Then, in section 4 and in order to complete the results obtained previously, we perform a non-parametric analysis based on the various instruments proposed by Quah (1996a, 1997). Subsequently, in section 5 we perform a preliminary study of the explanatory factors involved in regional mobility. The main conclusions are briefly presented in section 6.

2 Mobility as compensation for inequality

We will begin our analysis of mobility by investigating its role in compensating for inequality. Traditionally a high degree of mobility has been linked with lower long term inequality levels than tend to be detected in more reduced periods. One way of testing mobility, therefore, is to observe the relationship between cross-sectional and longitudinal inequality. Therefore, following common practice in the literature devoted to the dynamic analysis of personal income distribution, in this section we will examine the family of indices proposed by Shorrocks (1978a).

Let us consider a society with a population of H individuals, each of whom receives a given income over T periods, such that y_h^t denotes the income received by individual h in period t, where $h = 1, 2, \ldots, H$, and $t = 1, 2, \ldots, T$. If $\mu^t = \frac{1}{H}\sum_{h=1}^{H} y_h^t$ is the average income of the H individuals in period t, the average accumulated income over the T periods considered will be given by $\mu = \sum_{t=1}^{T} \mu^t$. Likewise, let Y be the vector of income accumulated by the H individuals over the T periods. That is, $Y = (Y_1, Y_2, ..., Y_H)$, where $Y_h = \sum_{t=1}^{T} y_h^t$. Finally, Y^t denotes the vector of the incomes of the H individuals in period t. That is, $Y^t = (y_1^t, y_2^t, ..., y_H^t)$.

We will now denote by $I(Y)$ the set of inequality measures that are convex functions of the the relative incomes. Then, given the convexity of the function, it can be written as:

$$I(Y) = h\left(\frac{Y}{\mu}\right) = h\left(\frac{\sum_{t=1}^{T} Y^t}{\mu}\right) = h\left(\sum_{t=1}^{T} \omega^t \frac{Y^t}{\mu^t}\right) \leq \sum_{t=1}^{T} \omega^t h\left(\frac{Y^t}{\mu^t}\right) \tag{1}$$

where ω^t is the ratio of average incomes in period t to the average accumulated income, such that $\omega^t = \frac{\mu^t}{\mu}$. Thus from expression (1) we have that:

$$I(Y) \leq \sum_{t=1}^{T} \omega^t I(Y^t) \tag{2}$$

That is, the inequality index of the incomes accumulated during the T periods considered can not exceed the weighted sum of the inequality indices for each of the individual periods. The rigidity index proposed by Shorrocks (1978a) is therefore defined as:

$$R(Y, Y^t) = \frac{I(Y)}{\sum_{t=1}^{T} \omega^t I(Y^t)} \tag{3}$$

with $R(Y, Y^t) \leq 1$. Note that the above expression is valid only for inequality measures that are convex functions of the relative incomes. This constraint does not impose a major drawback, however. Indeed, most of the indices commonly used (the Gini index, the family of Theil indices, Atkinson's indices, etc.) fufil this property[8].

The index $R(Y, Y^t)$ gives the value at which inequality diminishes as the time period considered is extended. Thus, for example, if $R(Y, Y^t) = 0.90$, income inequality over a given period will be 90 per cent of the average inequality corresponding to the set of subperiods contemplated. In other words, this index measures the stability of inequality as the sample period progresses. Indeed, if inequality remains stable as the period is extended, we will have:

$$I(Y) = \sum_{t=1}^{T} \omega^t I(Y^t) \tag{4}$$

with $\frac{Y^t}{\mu^t}$ independent of t, such that $R(Y, Y^t) = 1$. In other words, relative incomes will not vary at all over time, a feature that is characteristic of a completely immobile society. In a society with a certain degree of mobility, however, it is to be expected that there will be more frequent and wider variations in relative incomes, which would mean that the value of $R(Y, Y^t)$ would be less than one. Thus, $R(Y, Y^t) = 0$ would indicate a case of perfect mobility in which $I(Y) = 0$[9]. Therefore, $R(Y, Y^t)$ gives us the following measure of mobility:

$$RM(Y, Y^t) = 1 - \frac{I(Y)}{\sum_{t=1}^{T} \omega^t I(Y^t)} \tag{5}$$

In contrast to the literature devoted to the study of personal income distribution, however, we are concerned in this study with regions, each of which contains a variable set of individuals. We will therefore denote per capita income in region i over the period t by x_i^t, where $x_i^t = \frac{X_i^t}{N_i^t}$, and X_i^t and N_i^t are respectively the income and population of region i, $i = 1, 2, \ldots, n$. Likewise, let p_i^t be the relative frequency of region i in period t, $p_i^t = \frac{N_i^t}{N^t}$, with $N^t = \sum_{i=1}^{n} N_i^t$. The associated per capita income and population distributions will therefore be given by $x^t = (x_1^t, x_2^t, \ldots, x_n^t)$ and $p^t = (p_1^t, p_2^t, \ldots, p_n^t)$[10]. Finally, let us assume that $x^t \in \mathbb{R}_+^n$, while $p^t \in \mathbb{R}_{++}^n$.

Given, however, that our unit of reference is not the individual, we must consider the specific characteristics of regional mobility, where, over time, each region registers variations in per capita income, which, in turn, are known to be the result of changes in income and population. Thus, the evolution of the various inequality measures reflects variations both in per capita income and in the population share of each region. However, if we consider mobility as the capacity of regions to modify their relative positions in terms of development over time, we must focus our analysis exclusively on per capita income variations, and eliminate the impact of population shifts. To better comprehend this idea, let us consider the following example. Let us imagine that we have data for a period of several years on the regional per capita income distribution in a country with two regions. Further, let us suppose that the per capita incomes remain unaltered over time. However, a variable share of the population moves from one region to the other each year. In a situation such as this, Shorrocks' rigidity index would vary over time, as a consequence of the modification in the inequality indices in the different periods. Nevertheless, according to our chosen definition of mobility, we would in theory have to say that per capita income distribution in the country in question is completely immobile.

In order to overcome this problem, we will from now on consider that the population remains constant, taking as reference the average population over the time period considered. That is, $p_i^t = \bar{p}_i$, where $\bar{p}_i = \frac{1}{T}\sum_{t=1}^{T} p_i^t$. Likewise, for the n regions $\bar{p} = (\bar{p}_1, \bar{p}_2, ..., \bar{p}_n)$[11]. We will also use the n-dimensional vector $\hat{x}$ to denote aggregate per capita incomes over the T periods considered. Thus, $\hat{x} = (\hat{x}_1, \hat{x}_2, ..., \hat{x}_n)$, where $\hat{x}_i = \sum_{t=1}^{T} x_i^t$ is the aggregate per capita income of region i over the T periods.

From now on, therefore, we can define Shorrocks' rigidity index (1978a) adapted to the specific characteristics of regional mobility as

$$R^*(\hat{x}, x^t, \bar{p}) = \frac{I(\hat{x}, \bar{p})}{\sum_{t=1}^{T} \omega^t I(x^t, \bar{p})} \tag{6}$$

where $\omega^t = \frac{\mu^t}{\mu}$, and $\mu = \sum_{i=1}^{n} \bar{p}_i \hat{x}_i$.

Thus, the corresponding measure of mobility will be[12]:

$$RM^*(\hat{x}, x^t, \bar{p}) = 1 - \frac{I(\hat{x}, \bar{p})}{\sum_{t=1}^{T} \omega^t I(x^t, \bar{p})} \tag{7}$$

Figure 1 shows the results of the calculation of $RM^*(\hat{x}, x^t, \bar{p})$ for the EU regional distribution of per capita income between 1977 and 1999, taking different time periods ($m = 1, 2, \ldots, 23$). However, to check the sensitivity of the results to the choice of inequality index used to calculate $RM^*(\hat{x}, x^t, \bar{p})$, we have opted to incorporate into the analysis various measures of inequality, since each index features a different way of aggregating the information contained in the distribution[13]. Following this approach, we selected the following measures: the coefficient of variation, $CV(x)$, the family of Theil indices, $T(\beta)$ with $\beta = 0$ and $\beta = 1$, and the normative Atkinson index for different levels of inequality aversion, $A(\varepsilon)$ with $\varepsilon = 0.5$ and $\varepsilon = 2$.

[INSERT FIGURE 1 HERE]

The results obtained show values of the mobility measure based on Shorrocks' rigidity index (1978a) that increase gradually as the period of reference is extended, independently of the inequality measure that is used (note that the ordinate axis has a scale of 0 to 0.1). This reveals that regional inequality in Europe declines very slowly when longer time intervals are considered. Therefore, the influence of transient variability in regional disparities within the EU appears to be quite limited, so that most of the observed inequality in this respect can be considered permanent. To illustrate this, Figure A1 displays the $R^*(\hat{x}, x^t, \bar{p})$ indices for the whole period 1977-1999. According to these, depending on the inequality index used to calculate $R^*(\hat{x}, x^t, \bar{p})$, regional inequality in per capita income in the European context over the twenty-three years considered falls within the range of 93 to 98 per cent of average inequality for the set of subperiods contemplated. This suggests that, according to $RM^*(\hat{x}, x^t, \bar{p})$, regional per capita income distribution in the EU is quite rigid and, therefore, barely mobile.

Nevertheless, detailed analysis of the information supplied in Figure 1 enables us to observe that the results obtained differ slightly according to the inequality index used in the calculation of $RM^*(\hat{x}, x^t, \bar{p})$. Both Theil indices follow a similar trend, though there appears to be a slight reduction in mobility as β increases. It is worth recalling, in this respect, that the β parameter captures the sensitivity of $T(\beta)$ to transfers between individuals at different points in the distribution. Following Shorrocks (1980), it can actually be shown that, as β increases, $T(\beta)$ becomes more sensitive to transfers in the upper end of the distribution. Also, as might be expected from the above results, mobility becomes greater as the value of ε increases. In fact, as is known, the higher the value of the inequality aversion parameter, the greater the sensitivity of Atkinson's index to what happens in the lower end of the distribution. The empirical evidence presented so far, therefore, appears to suggest that the reduction in inequality that takes place as the time interval is extended is greater in the European regions with lower per capita income levels.

3 Regional mobility: an analysis based on transition matrices

The measure of mobility considered in the previous section may in certain circumstances present some drawbacks relating to the significance of changes in the relative positions of the regions according to per capita income. To illustrate this problem, let us consider another example that highlights the multidimensional nature of mobility. Let us imagine a country with two regions, one of which enjoys some comparative advantage over the other, in terms, say, of its spatial location. In a situation of this kind, the region in question will, *ceteris paribus*, systematically register higher growth rates, giving rise to an increase in regional disparities, even after an initial situation of hypothetical equality. In other words, the rank ordering of the two regions will remain unaltered over time. In a context such as this, $RM^*(\hat{x}, x^t, \bar{p})$ will present positive values, though it could be argued that there is no mobility in the regional income distribution.

Keeping this fact in mind, in this section we have considered a new approach to the analysis of intradistributional mobility, based on the observation of changes experienced by relative positions of the various regions.

One of the most intuitive options when approaching mobility studies in this way is to construct transition matrices. In order to define the concept of transition matrix, let us now suppose that we have classified the different regions in the distribution into m exhaustive and mutually exclusive classes according to their per capita income level. Further, let us imagine that we have information on the distribution of interest for two moments in time, t_0 and t_1. In a case such as this, the matrix that summarizes the probabilities of regions shifting from one class to another between t_0 and t_1 is known as a transition matrix. Supposing, therefore, that the probabilities can be reasonably estimated from the corresponding relative frequencies, the transition matrix associated with the transformation experienced by the distribution between t_0 and t_1 ($x^{t_0} \longrightarrow x^{t_1}$), will be the square matrix $\Pi(x^{t_0}, x^{t_1}) = \left[\pi_{jk}(x^{t_0}, x^{t_1})\right] \in \mathbb{R}_+^{m \times m}$, where $\pi_{jk}(x^{t_0}, x^{t_1})$ denotes the proportion of regions that belonged to class j in t_0 and have shifted to class k in t_1. According to this definition, we have that $\sum_{k=1}^{m} \pi_{jk}(x^{t_0}, x^{t_1}) = 1$ for any

$j = 1, 2, \ldots, m$, so that $\Pi(x^{t_0}, x^{t_1})$ is a stochastic matrix.

The literature devoted to the dynamic study of personal income distribution have designed numerous measures of mobility based on transition matrices[14]. From this wide range of options we began by considering the following index based on Shorrocks (1978b)[15]:

$$SM^*(\Pi, \rho) = \frac{1 - \sum_{j=1}^{m} \rho_j \pi_{jj}}{1 - \frac{1}{m}} \tag{8}$$

where ρ_j denotes the population share in relation to the total of class j. That is, $\rho_j = \frac{N_j}{N}$[16].

This measure captures those aspects of the mobility concept that refer to the independence with regard to the initial situation. Nevertheless, $SM^*(\Pi, \rho)$ is of limited validity if the aim is to highlight that dimension of mobility that is related to movement *per se*[17], since it is calculated exclusively from those elements that form the main diagonal of the transition matrix, thereby ignoring the rest of the elements in Π. To overcome this problem associated with the use of $SM^*(\Pi, \rho)$, we opted to consider in addition the following index proposed by Bartholomew (1973):

$$BM^*(\Pi, \rho) = \sum_{j=1}^{m} \sum_{k=1}^{m} \rho_j \pi_{jk} |j - k| \tag{9}$$

The next step is to select an appropriate definition for each of the various classes. Faced with this problem, we decided to adopt a solution that enables us to obtain reasonably accurate information on regional movements across a sufficiently large number of groups, without risking any loss of representativity of the results. Thus, we divided the regions that make up the distribution under analysis into five exhaustive and mutually exclusive classes, according to their per capita income in relation to the European average, which was assigned a value of 100: [0,75), [75,90), [90,110), [110,125) and [125,+∞)[18].

[INSERT FIGURE 2 HERE]

Figure 2 shows the calculations of $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ after estimating the corresponding transition matrices. In addition, in order to isolate the effect of transient per capita income fluctuations associated with annual changes, we opted to use in our analysis time periods of different length, thus we were also able to distinguish between short and medium term mobility.

The results obtained reveal that regional per capita income distribution exhibits greater mobility, the longer the time interval taken as a reference. Thus on average, 91 per cent of the regions considered continued in the same class after a year. Taking the period as a whole, however, the percentage drops to 63 per cent.

It is also worth stressing that the two mobility indices considered follow very similar trends. Given that the main difference between them lies in the different valuation given to shifts between classes, this result suggests a relatively low degree of intradistributional mobility[19]. Further confirmation of this is to be found in the various transition matrices estimated, which exhibit the highest values around the main diagonal[20].

Whatever index is used, the empirical evidence presented shows a reduction in the mobility of the EU regional per capita income distribution between 1997 and 1999. Nevertheless, since mobility did not fall at an even rate over time, it is possible to identify a series of separate stages each with its distinguishing features. Thus, the main reduction in $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ took place between 1977 and the early eighties. From then onwards, however, there is a change of trend leading to an increase in regional mobility continuing until the end of that decade. During the early nineties, there was a further decrease in regional mobility, which, however, seemed to mark the beginning of a new stage, characterized by a new rise in $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$[21].

In this context, however, it is necessary to stress that the above results cannot be valued normatively without taking into account the degree of inequality observed in the distribution under analysis. In this respect, a large number of studies have coincided in reporting a lack of regional convergence in per capita income in the European context from the mid-seventies onwards [Armstrong (1995), Neven and Gouyette (1995), López-Bazo *et al.* (1999), Rodríguez-Pose (1999), etc.]. The analysis performed in this section, for its part, shows that this maintenance of territorial imbalances has coincided in time with a process of consolidation in the relative positions of the various regions, which stresses the need for an active regional policy at European level [22].

Finally, in light of the volatility of $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ in short term observations, we performed a preliminary analysis of the relationship between the economic cycle and regional mobility trends in the European context. To this end we estimated the statistical correlation between per capita income growth rates in the EU and annual fluctuations in the two mobility measures considered in this section. We then repeated the exercise incorporating the assumption that economic cycle influences on regional mobility with a lag[23]. In both cases, however, the correlation coefficients, though positive, were not statistically significant[24].

4 A non-parametric analysis of intradistributional mobility

By means of the various tools employed in the preceding section, we have explored the level and evolution of regional mobility in the EU between 1977 and 1999. It is necessary to bear in mind, however, that $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ were calculated on the basis of the information supplied by various transition matrices, obtained by dividing the distribution of interest into a series of exhaustive and mutually exclusive classes. However, since there is no procedure for finding the optimal number of classes in each case, the researcher is obliged to make an arbitrary decision in this respect[25].

To address this problem, Quah (1996a, 1997) suggests substituting the transition matrix with a stochastic kernel that reflects the probabilities of transition between a hypothetically infinite number of classes, reducing their size infinitesimally[26]. The stochastic kernel can be reached by estimating the density function of the distribution over a given period, $t + k$, conditioned by the values of a previous period, t. Specifically, the joint density function of the distribution at moments t and $t + k$ is estimated non-parametrically and normalized by the implicit marginal distribution at t in order to obtain the corresponding conditional probabilities.

[INSERT FIGURE 3 HERE]

Figure 3 shows the stochastic kernel estimated for the European regional per capita income distribution over a period of twenty-three years ($t = 1977$ and $t + k = 1999$)[27]. This three-dimensional graph informs about the probabilities associated with each pair of values in the first and last years of the study period. In other words, the stochastic kernel provides, in a way analogous to that of a discrete transition matrix, the probability distribution of 1999 per capita income for regions with a given per capita income in 1977. The peaks on the graph represent high levels of probability. Thus, if the probability mass is concentrated around the main diagonal, the intradistributional dynamics are characterized by a high level of persistence in the relative positions of the regions over time and, therefore, low mobility. If, on the other hand, the density is located mainly on the opposite diagonal to the main diagonal, this would indicate that regions at each end of the distribution exchange their relative positions throughout the period. Finally, the probability mass could, in theory, accumulate parallel to the t axis. This would reflect the existence of a convergence process in regional per capita incomes. In order to aid interpretation of the graph, Figure 3 also includes a contour plot on which the lines connect points at the same height on the three-dimensional kernel.

The results obtained fully uphold the conclusions reached in the previous analysis based on the data from the discrete transition matrices. Indeed, as can be seen from Figure 3, the mass of probability is concentrated around the main diagonal. As we are already aware, this shows that there was little mobility in the distribution of regional per capita income between 1977 and 1999. There is a general tendency, therefore, for the European regions to maintain their relative positions throughout the twenty-three years contemplated. By means of these tools we are also able to detect the fact that mobility patterns vary in terms of economic development levels. It is possible to observe, for example, how regions with a per capita income close to the European average exhibit a relatively higher degree of mobility over time, while those located at each end of the distribution are characterized by a stronger persistence in their relative positions. Indeed, the information provided by Figure 3 in this respect confirms that there is comparatively less mobility among more highly developed regions than among regions with low levels of per capita income over the time period considered[28].

In light of these results, we completed the above analysis with further information relating to the behavior of the regions situated at each end of the distribution under study, taking these to the ones in which per capita income fell outside the interval of 50 per cent to 150 per cent of the European average. Our calculations revealed that 27 per cent of the regions with a capita income below 50 per cent of the European average in 1977, continued in the same situation in in 1999. In fact, of the 22 regions whose per capita income in 1977 was below 50 per cent of the European average, only the Portuguese regions of Norte, Centro, Alentejo, Algarve, Azores and Madeira remained in the same situation twenty-three years later. However, out of the other 16, only the Spanish regions of Aragón, Baleares, Madrid, Cataluña and La Rioja had succeeded in raising their per capita income above 75 per cent of the European average, which is further support for the results obtained earlier. There is a different situation at the upper end of the distribution, however, where out of the 13 regions who began the period with a per capita income above 150 per cent of the European average, only the Swedish regions of Norra Mellansverige, Mellersta Norrland and Ovre Norrland, together with Valle d'Aosta and Groningen had dropped from that level by 1999, though none of them had fallen below 125 per cent of the European average.

5 Some explanatory factors for regional mobility

To round off the results obtained in the previous sections, we will now investigate the role played by a series of factors in accounting for the observed level of intradistributional mobility in the EU from 1977 to 1999. Our specific aim will be to ascertain why some regions have improved their relative position, while others have worsen over the twenty-three years considered.

Thus, our first step will be to determine which dependent variable to use in the analysis. If, for the study period considered, we wish to use data deriving from one of the various mobility measures calculated in the preceding pages, we will have, at best, only twenty-two values for each index. Needless to say, even if we were willing to consider only interannual mobility, such a degree of freedom would be clearly insufficient for the analysis to be statistically significant. To address the problems surrounding this issue, we opted for the alternative of considering an individual measure of regional mobility, $\Delta RNK_i(t_0, t_1)$, which assigns to each region its shift in the rank ordering in terms of per capita income over a given period. Under these conditions, it is worth noting that any upward shift in the ranking on the part of one region inevitably means a downward shift of the same magnitude for other regions. That is, $\sum_{i=1}^{n} \Delta RNK_i(t_0, t_1) = 0$. Certainly, the use of $\Delta RNK_i(t_0, t_1)$ will involve some drawbacks that will need to be borne in mind when it comes to making an accurate interpretation of the results of the empirical analysis. The most obvious of these is the fact that this indicator only registers levels of mobility that bring about a change in the ranking of the regions. In other words, if throughout the course of the time period considered there are no changes in per capita income sufficient to cause an alteration in the ranking, $\Delta RNK_i(t_0, t_1)$ will take a null value for any $i = 1, 2, ..., n$, in spite of any movement that might have taken place in the distribution. Unlike standard mobility measures, however, $\Delta RNK_i(t_0, t_1)$ provides information about the direction of regional shifts, so that it is possible to tell which regions have risen and which have fallen in the ranking over time. Likewise, as pointed out earlier, the use of this indicator will increase the robustness of the subsequent analysis, by addressing the problems arising from the lack of degrees of freedom.

Having established the dependent variable, we then investigated to determine how far the initial per capita income level ($GVApc_{i0}$) contributes to explain observed regional mobility in the European context. Moreover, the importance of the role of the sectoral composition of economic activity in regional growth processes is widely known[29]. It is therefore reasonable to assume that the initial productive structure and the changes that have taken place therein over the course of time may be related to shifts in the regional ranking in terms of per capita income. Taking into account this idea, we decided to introduce into our model the initial share in regional employment of agriculture (EAG_{i0}), the financial sector (EFS_{i0}), and the non-market services ($ENMS_{i0}$); together with the variation in these variables over the period analyzed (ΔEAG_i, ΔEFS_i and $\Delta ENMS_i$). It is in fact common practice in the literature devoted to the estimation of convergence equations to include a variable to capture the size of the agricultural sector, in order to control for differences in the sectoral composition of activity across the different territorial units to be analyzed[30]. However, bearing in mind the process of increasing tertiarization that has been taking place in the European economy for the last few decades[31], we decided to consider, in addition, the role played in this context by advanced services and public employment, which we approximated, respectively, with EFS_{i0}, ΔEFS_i, $ENMS_{i0}$ and $\Delta ENMS_i$.

Thus, our proposed model to explain the regional mobility registered in the EU between 1977 and 1999 is defined as follows:

$$\begin{aligned}\Delta RNK_i &= \beta_0 + \beta_1 GVApc_{i0} + \beta_2 EAG_{i0} + \beta_3 \Delta EAG_i + \beta_4 EFS_{i0} + \\ &\quad + \beta_5 \Delta EFS_i + \beta_6 ENMS_{i0} + \beta_7 \Delta ENMS_i + u_i \qquad (10)\end{aligned}$$

where u_i is the corresponding error term.

Table 1 shows the estimation of the above model by ordinary least squares (OLS) for different time periods. Before interpreting the results, however, it should be borne in mind that several studies have underlined the relevance of the spatial dimension in explaining observed territorial imbalances in the EU[32]. The analyses carried out in these studies suggest, in particular, the possible presence of some type of geographical externality in the European context, in as far as spatially close regions tend to enjoy similar levels of development.

[INSERT TABLE 1 HERE]

In order to assess the importance of this issue within the context of this paper, we defined a spatial weighting matrix, W, which allows to capture the strength of the interdependence between each pair of regions i and j. For this, a first option is to use the notion of physical contiguity of first order, according to which $w_{ij} = 1$ if regions i and j are geographically adjacent and 0 otherwise[33]. However, in order to take into account the direct interaction of all the regions considered, we decided instead to use a spatial weighting matrix standardized by rows based on the inverse square distance among the centroid of the different regions[34].

We then proceeded to calculate the Moran's I and various Lagrange multiplier tests using the residuals provided by the OLS estimations (Burridge, 1980; Anselin, 1988; Anselin *et al.*, 1996). The results obtained indicate the existence of a specification problem in the model considered suggesting, in accordance with Anselin and Florax (1995), the need to include a spatial lag of the dependent variable (spatial lag model). This fact implies that the OLS estimations will be biased and inconsistent. In order to overcome this problem, we estimated the following model by maximum likelihood (ML-LAG):

$$\begin{aligned}\Delta RNK_i &= \beta_0 + \beta_1 GVApc_{i0} + \beta_2 EAG_{i0} + \beta_3 \Delta EAG_i + \beta_4 EFS_{i0} + \\ &\quad + \beta_5 \Delta EFS_i + \beta_6 ENMS_{i0} + \beta_7 \Delta ENMS_i + \beta_8 W \Delta RNK_i + u_i \quad (11)\end{aligned}$$

As Table 1 shows, the results obtained reveal an inverse relationship between ΔRNK_i and the initial per capita income level, which allows us to complete and qualify some of the findings from the analysis performed in the preceding section. It is also worth noting the low dynamism of the agricultural regions. Indeed, the presence in 1977 of a relatively important agricultural sector or the growth of this sector in employment terms, are found to be associated with downward shifts in the regional ranking. Meanwhile, EFS_{i0} is also statistically significant. This suggests that upward shifts in the regional ranking are linked to the share in the economy of certain types of advanced services of high productivity. In any event, the increase in non-market services is negatively correlated with ΔRNK_i. This is consistent with the empirical evidence presented by Rodríguez-Pose and Fratesi (2004a), who stress the fact that the European peripheral regions of Europe characterized by high levels of public employment presented more moderate growth rates than the rest between 1980 and 2000.

Next, in order to detect possible variations in behavior patterns over time, we decided to repeat the analysis for various time intervals of a shorter duration. However, the results for the 1977-1988 subperiod are very similar to those just discussed for the period as a whole. In particular, in this case, the only difference arises from the fact that the increase in employment in the financial sector appears to have a negative effect on the dependent variable.

When analyzing the 1988-1999 subperiod, we introduced a slight modification to the model we had been estimating so far, in order to obtain a first impression of the relationship between the EU regional policy and observed intradistributional mobility. This involved the introduction of a dummy variable, $RO1_i$, to enable us to identify all the regions that held Objective 1 status in any of the various programming periods[35]. In this way we will be able to see whether regions that have benefited from priority treatment under EU regional policy perform differently from the rest. In this respect, the information provided by Table 1 suggests that $RO1_i$ is unrelated to variations in the dependent variable. This result should nevertheless be interpreted with caution. On the one hand, it should be borne in mind that Boldrin and Canova (2001) and Rodríguez-Pose and Fratesi (2004b) both insist on the low mobility of the less developed regions of the EU during the nineties[36]. It would be extremely risky, however, to judge something as complex as the relationship between EU regional policy and the dynamics exhibited over the course of the last decade by the Objective 1 regions exclusively on the basis of the results of an analysis of this nature.

Finally, with respect to the rest of the explanatory variables considered in this study, the main difference between the estimations for the 1988-1999 subperiod and those for the period as a whole relates to the fact that during those 12 years EFS_{i0} is not statistically significant. In other words, it is not possible to establish any link between the shifts that have taken place in the regional ranking and the share of the financial sector in total employment in 1988.

6 Conclusions

In this article we have examined mobility in the regional distribution of per capita income in the EU between 1977 and 1999 from several complementary perspectives. We began by calculating a wide range of measures based on the literature devoted to the dynamic analysis of personal income distribution. Our results show a decrease in mobility within the distribution under study over the period of observation. A further feature of note is relatively low level of intradistributional mobility. This conclusion is in fact confirmed when stochastic kernel and contour plot are estimated for a series time intervals of different length. Therefore, with only a few exceptions, the European regions tended to maintain their relative positions in the ranking over the twenty-three years considered. All of this underlines the need for the EU to reinforce its regional development policies.

Our results also show that regional mobility patterns vary as a function of economic development. In fact, the regions with a per capita income close to the European average tended to register a relatively higher mobility degree over time, while those at either end of the distribution were characeterized by a stronger persistence in their relative positions. However, less developed regions showed greater mobility than regions located at the upper end of the distribution.

Finally, we carried out a regression analysis by means of spatial econometric techniques in order to identify some of the explanatory causes of regional mobility in the EU. The results obtained for the 1977-1999 period reveal the existence of an inverse relationship between upward shifts in the regional ranking and initial per capita income levels. Furthermore, the presence at the onset of the period of a relatively large agricultural sector or the increase in the share of employment in this sector are found to be associated with downward shifts in the regional relative positions. In fact, an increase in employment in non-market services has a similar effect, in contrast with what occurs with the financial sector. Finally, according to our analysis, the Objective 1 regions failed in general terms to improve their relative positions over the 1988-1999 period, in spite of the priority treatment they were given under EU regional policy.

Notes

[1]A review of the main results can be found in Armstrong (2002) or Terrasi (2002).

[2]Article 2 of the Treaty of the EU specifically states that "The Community shall have as its task (...), to promote (...) a harmonious, balanced and sustainable development of economic activities, (...) sustainable and non-inflationary growth, (...) a high degree of competitiveness and convergence of economic performance (...).".

[3]The properties of symmetry and scale independence do not constitute a major limitation. Indeed both are basic properties that any inequality index can be reasonably expected to fulfill (Cowell, 1995). In any event, for the purposes of our example, we can overcome the need for the inequality index to satisfy the property of scale independence by simply assuming the per capita incomes of A and B to be equal.

[4]Two exceptions worth mentioning are the contributions made by López-Bazo *et al.* (1999) and Cuadrado *et al.* (2002).

[5]Save for a few exceptions, the recent literature on convergence does not take into account differences in population across territorial units, and uses almost exclusively unweighted statistics. See, for example, Salas (2002) or Goerlich (2003).

[6]Indeed, as stated in Fields and Ok (1999), considerably different approaches are currently taken in the study of inequality and mobility. Nevertheless, over the course of the last decade, major theoretical advances have been made in the analysis of intradistributional mobility. In particular, there have been proposed a series of measuring procedures with similar axiomatic contents to those used in the study of inequality.

[7]Lack of complete series, however, has obliged us to eliminate from the analysis the member States

newly admitted to the European in May 2004, the *Länder* of former East Germany, The French overseas departments and the Spanish territories in North Africa. Nevertheless, the appendix includes a complete list of all the regions considered in this study.

[8]The most outstanding exception is the variance of log of incomes.

[9]If $I(Y) = 0$, we have that $Y_1 = Y_2 = \ldots = Y_H$.

[10]Obviously, $\sum_{i=1}^{n} p_i^t = 1$.

[11]Again, $\sum_{i=1}^{n} \bar{p}_i = 1$.

[12]Note that in the previous example $R^*(\hat{x}, x^t, \bar{p}) = 1$, therefore $RM^*(\hat{x}, x^t, \bar{p}) = 0$.

[13]For further details relating to this issue, see Chakravarty (1990) or Cowell (1995).

[14]In relation to this, see, for example, Prais (1955), Bartholomew (1973), Bibby (1975), Shorrocks (1978b), Sommers and Conlisk (1978) or Conlisk (1985, 1990).

[15]The mobility measure proposed by Shorrocks (1978b) is given by:

$$SM(\Pi) = \frac{m - tr(\Pi)}{m - 1}$$

where $tr(\Pi)$ denotes the trace of the matrix Π. Note that, in contrast to what occurs with $SM^*(\Pi, \rho)$, this index assigns identical weight to each of the m classes. Indeed, if $\rho_j = \frac{1}{m}$ for any $j = 1, 2, \ldots m$, it is obtained that $SM^*(\Pi) = SM(\Pi)$.

[16]Given that matrix Π is stochastic and $N_i > 0$ for any $i = 1, 2, \ldots, n$, then $\rho_j > 0$ for any $j = 1, 2, \ldots, m$.

[17]For further details regarding this issue, see Fields and Ok (1999).

[18]This classification was adopted, for example, by López-Bazo *et al.* (1999) or Cuadrado *et al.* (2002).

[19]Neven and Gouyette (1995) and López-Bazo *et al.* (1999) reach a similar conclusion for a more reduced geographical area and a shorter time period than considered in this article.

[20]The medium and full term transition matrices are included in the appendix. The rest, which are not shown for lack of space, are available from the authors upon request.

[21]In order to test the robustness of the above results, we recalculated $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ for eight-category classification of the European regions, based on the following per capita income levels: [0,50), [50,75), [75,90), [90,100), [100,110), [110,125), [125,150) and [150,+∞). The results, which are shown in the appendix, are very similar to those we have just discussed.

[22]Note that, for a given level of inequality, high mobility would be a sign of strong cyclical variability in regional incomes. In this kind of context, regional policy should address the need to mitigate adverse cyclical effects before applying traditional convergence policies.

[23]In relation to this, see Fischer and Nijkamp (1987).

[24]Quah (1996b) obtains a similar finding for the United States.

[25]In relation to this question, see Kremer *et al.* (2001).

[26]See Stockey and Lucas (1989).

[27]Gaussian kernel functions are used in all cases, while the optimal smoothing parameter values have been selected following Silverman (1986, p. 47).

[28]In order to test the robustness of the results, we decided to repeat the above analysis using data only for the subperiods 1977-1988 and 1988-1999. The results, shown in the appendix, are very similar to those discussed in this section.

[29]With reference to the European case, see, for example, the works of Paci (1997) and Gil *et al.* (2002).

[30]Interested readers will find a review of the main results obtained in this type of studies in Magrini (2004).

[31]See European Commission (1999).

[32]In relation to this, see Fingleton and McCombie (1999), López-Bazo *et al.* (1999, 2004) or Le Gallo and Ertur (2003).

[33]This is in fact the option chosen by López-Bazo *et al.* (1999) or Rey and Montouri (1999) among others.

[34]It should be noted in this respect that the use of a matrix of this nature is consistent with the arguments employed to support gravitational models. For further details in relation to this issue, see Anselin (1996) and Anselin and Bera (1998).

[35]Let us not forget that the Objective 1 regions became a key element in EU regional policy after the Structural Fund reform in 1988.

[36]There are obviously some exceptions to this general trend. This is the case of Southern and Eastern Ireland or the Abruzzi in Italia, for example.

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Appendix

[INSERT FIGURE A1 HERE]

[INSERT FIGURE A2 HERE]

[INSERT FIGURE A3 HERE]

[INSERT FIGURE A4 HERE]

[INSERT TABLE A1 HERE]

[INSERT TABLE A2 HERE]

List of Figures and Tables

Figure 1: $RM^*(\hat{x}, x^t, \bar{p})$ index for various inequality measures.

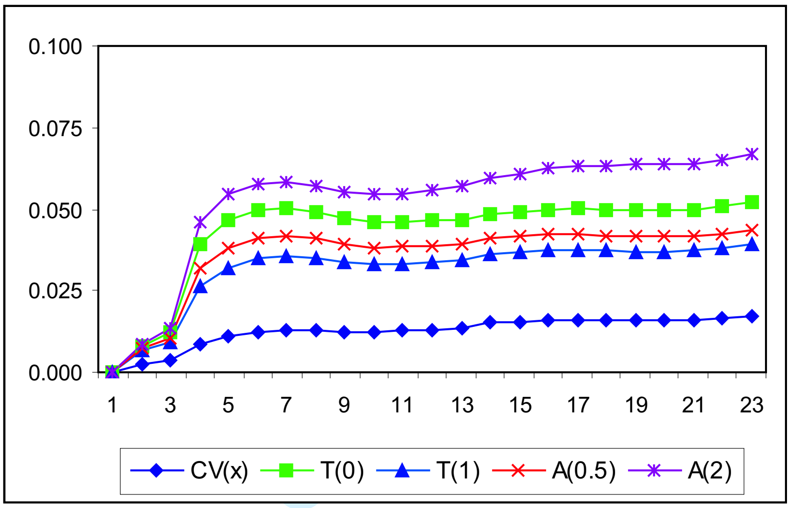
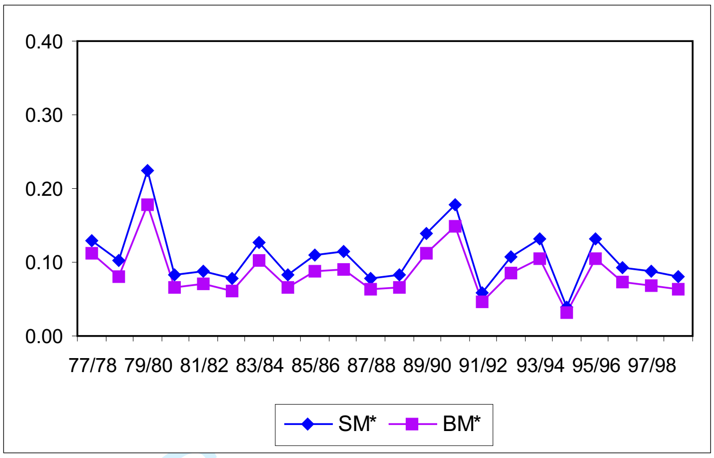
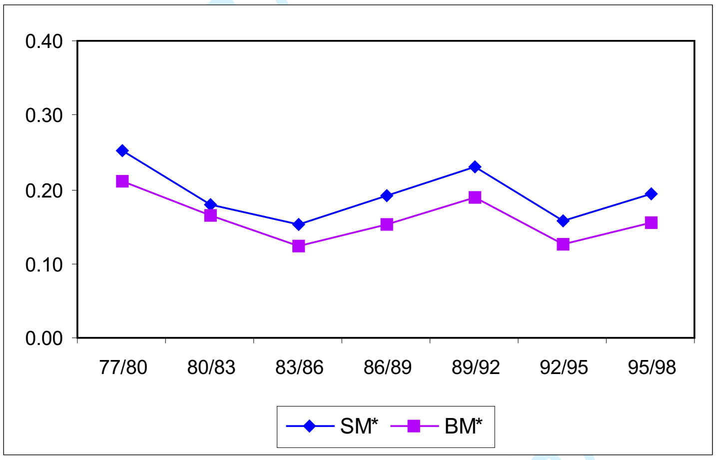


Figure 2: Regional mobility measured by $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$, $m = 5$.





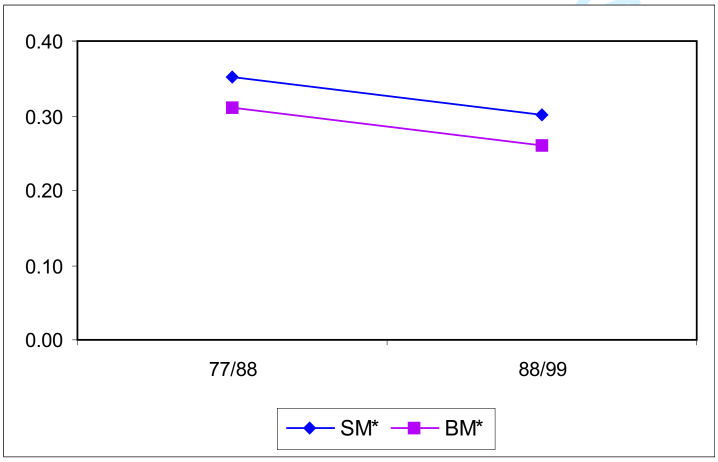


Figure 3: Stochastic Kernel and contour plot of regional per capita income distribution, 1977-1999.

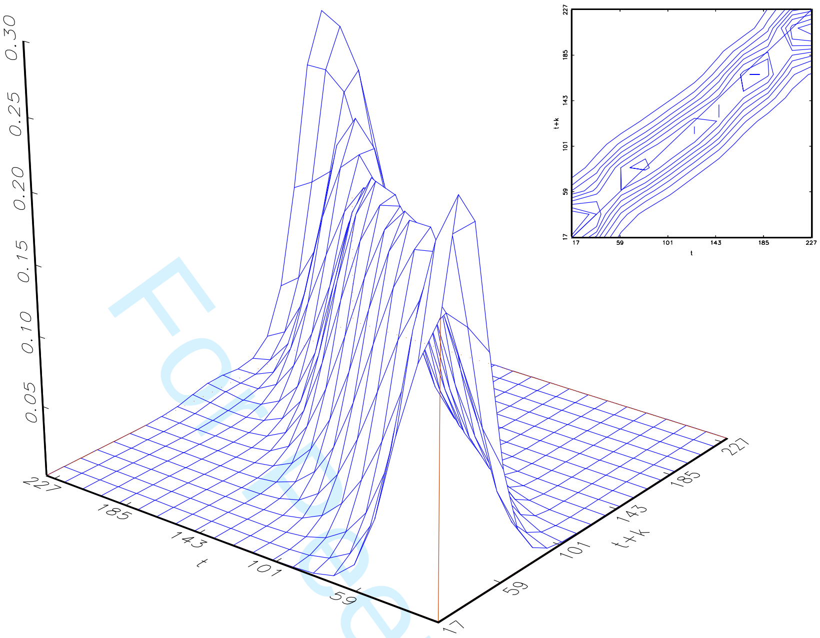


Figure A1: $R^*(\hat{x}, x^t, \bar{p})$ index, 1977-1999.

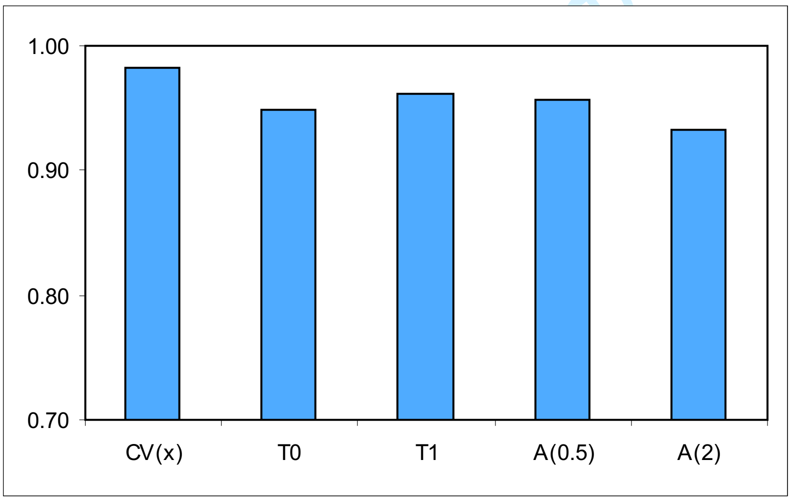
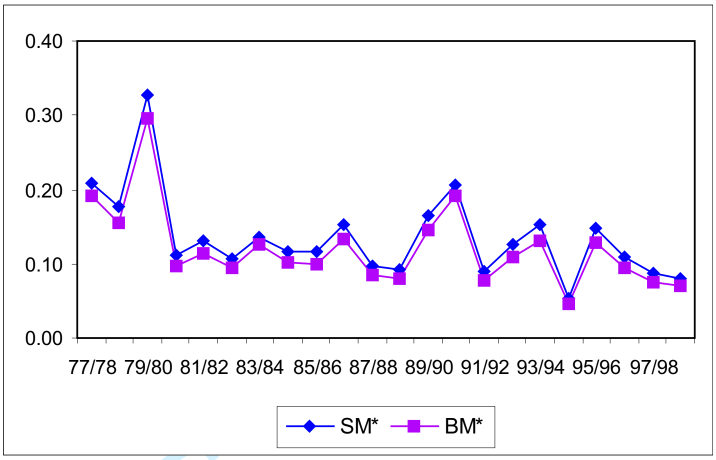
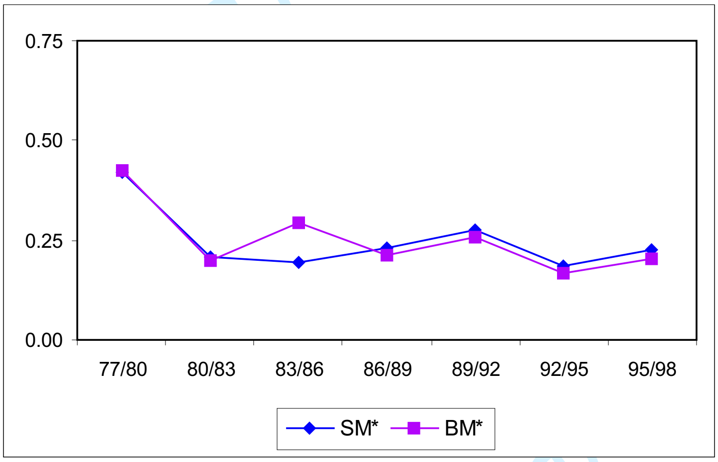


Figure A2: Regional mobility measured by $SM^{*}(\Pi,\rho)$ y $BM^{*}(\Pi,\rho)$, $m=8$.





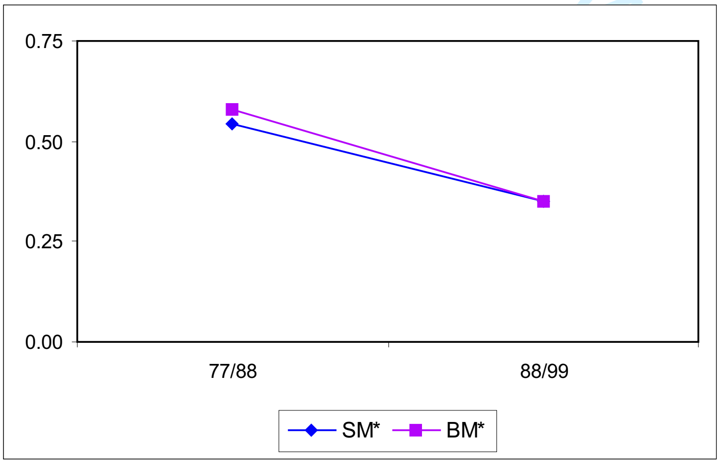


Figure A3: Stochastic kernel and contour plot of regional per capita income distribution, 1977-1988.

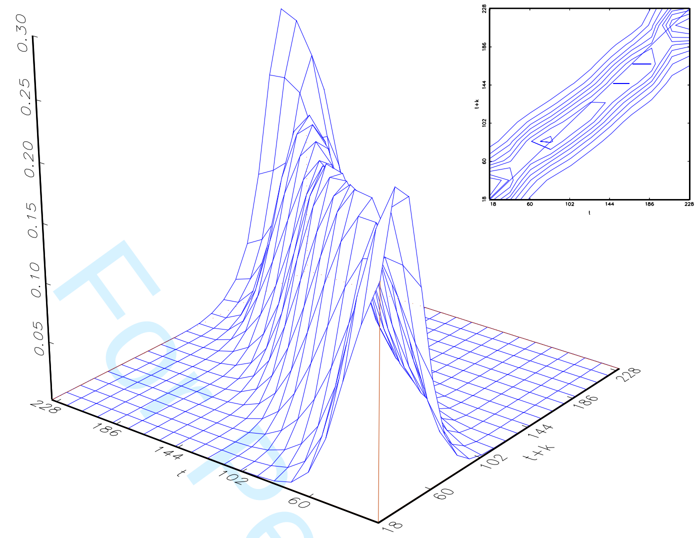


Figure A4: Stochastic kernel and contour plot of regional per capita income distribution, 1988-1999.

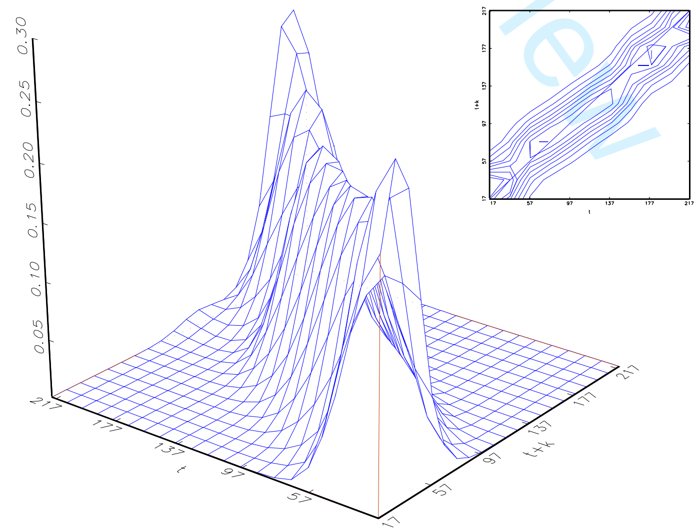


Table 1: Explaining factors of regional mobility.

	1977-1999		1977-1988		1988-1999	
Variable	OLS	ML-LAG	OLS	ML-LAG	OLS	ML-LAG
$Constant$	0.3225	0.3104	0.2125	0.2180	0.0947	0.1234
	(0.035)	(0.026)	(0.062)	(0.041)	(0.313)	(0.152)
$GVApc_{i0}$	-0.3753	-0.3335	-0.2855	-0.2471	-0.0607	-0.0877
	(0.000)	(0.000)	(0.000)	(0.000)	(0.041)	(0.001)
EAG_{i0}	-0.1784	-0.1601	-0.0947	-0.0925	-0.0419	-0.0326
	(0.000)	(0.000)	(0.000)	(0.000)	(0.047)	(0.089)
ΔEAG_i	-0.3382	-0.2746	-0.3104	-0.2802	-0.1262	-0.0861
	(0.000)	(0.000)	(0.000)	(0.000)	(0.006)	(0.035)
EFS_{i0}	0.5644	0.4063	0.7308	0.5425	-0.0229	0.0412
	(0.012)	(0.059)	(0.000)	(0.003)	(0.856)	(0.727)
ΔEFS_i	-0.5833	-0.3214	-0.8135	-0.6274	0.9885	0.6528
	(0.236)	(0.468)	(0.019)	(0.064)	(0.046)	(0.157)
$ENMS_{i0}$	0.0415	0.0357	0.0355	0.0289	0.0105	0.0054
	(0.322)	(0.358)	(0.297)	(0.367)	(0.638)	(0.790)
$\Delta ENMS_i$	-0.2098	-0.1655	-0.2367	-0.2235	-0.1667	-0.1088
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.004)
RO_i					-0.2904	-0.4351
					(0.529)	(0.302)
$W\Delta RNK_i$		0.7039		0.5786		0.6704
		(0.000)		(0.000)		(0.000)
$\overline{R}^2$	0.2634	0.3985	0.3229	0.3987	0.1450	0.2784
Log L		-905.65		-847.33		-808.15
I-Moran	6.882		4.700		6.186	
	(0.000)		(0.000)		(0.000)	
LMERR	35.234		15.644		27.392	
	(0.000)		(0.000)		(0.000)	
R-LMERR	0.054		0.317		0.006	
	(0.817)		(0.574)		(0.940)	
LMLAG	41.860		22.951		35.552	
	(0.000)		(0.000)		(0.000)	
R-LMLAG	6.679		7.624		5.166	
	(0.010)		(0.006)		(0.023)	

Notes: Numbers in parentheses are the corresponding p-values. Log L is the value of the log-likelihood function. LMERR (LMLAG) refers to the the Lagrange multiplier test used to examine the null hypothesis of no residual spatial autocorrelation versus an alternative autoregressive spatial error model (a spatial lag of the dependent variable), where R-LMERR (R-LMLAG) is its robust version. The standard errors were estimated from the variance-covariance matrix using the method proposed by White (1980).

Table A1: Transition matrix, 1977-1988.

Regions	ρ_j	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
46	0.19	0.81	0.17	0.02	0.00	0.00
45	0.20	0.18	0.67	0.15	0.00	0.00
46	0.20	0.07	0.13	0.71	0.07	0.02
24	0.15	0.00	0.00	0.33	0.63	0.04
36	0.26	0.00	0.00	0.03	0.22	0.75

Table A2: Transition matrix, 1988-1999.

Regions	ρ_j	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
48	0.22	0.98	0.00	0.02	0.00	0.00
44	0.19	0.15	0.55	0.30	0.00	0.00
50	0.23	0.00	0.16	0.68	0.16	0.00
26	0.16	0.00	0.00	0.15	0.73	0.12
29	0.20	0.00	0.00	0.07	0.10	0.83

Table A3: Transition matrix, 1977-1999.

Regions	ρ_j	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
46	0.19	0.78	0.09	0.13	0.00	0.00
45	0.20	0.33	0.43	0.24	0.00	0.00
46	0.20	0.07	0.17	0.59	0.15	0.02
24	0.15	0.00	0.04	0.21	0.71	0.04
36	0.26	0.00	0.00	0.14	0.17	0.69

Regional mobility in the European Union*

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Abstract

This paper examines regional mobility in the spatial distribution of per capita income in the European Union over the period 1977-1999. The methodology used to investigate this issue combines a series of measures taken from the literature devoted to the dynamic study of personal income distribution with a non-parametric analysis. The results show limited mobility in the distribution considered, and a decline in mobility over time. The empirical evidence presented indicates, moreover, that mobility patterns vary as a function of regional development levels. Additionally, the analysis carried out investigates the role played in explaining intra-distribution mobility by variables such as per capita income, population density, per capita expenditure in investment, market potential, and the share in total employment of agriculture, advanced services and non-market services.

Key words: Mobility, per capita income, regions, European Union.

JEL Code: D30, R11, R12.

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1 Introduction

In recent years, the issue of territorial imbalances in the European Union (EU) has been examined in numerous studies from a variety of different approaches (see Armstrong (2002) or Terrasi (2002) for a review of this literature). There are various reasons for the amount of interest surrounding this question. Among them is the fact that economic growth theory has advanced greatly over the last two decades, coinciding with the introduction of endogenous growth models in the mid eighties. Another, the need to reduce disparities in terms of development levels across the various European regions, is directly related to some of the basic principles behind the forming of the Union, especially since the introduction of the Single Act and the Maastricht agreements. In particular, one of the specific assumptions of the European integration program is that it will drive the growth of all Member States, thereby increasing economic and social cohesion[1].

Most analyses of regional per capita income disparities in Europe apply the concepts of sigma convergence and beta convergence introduced by Barro and Sala-i-Martin (1991, 1992), combining the information provided by various dispersion statistics with the estimation of convergence equations. However, as Quah (1993, 1996a,b; 1997) has repeatedly pointed out, not only does this approach raise a number of econometric problems, it also fails to capture a series of potentially interesting issues relating to the dynamics of the distribution in question. In particular, this type of analysis does not consider the possibility of regions modifying their relative positions over time, and thereby neglects the whole issue of intra-distribution mobility.

To illustrate the relevance of issues relating to the analysis of distribution dynamics, let us consider the following example. Let us assume that we have regional income and population data for several years for two countries, A and B, each of which is in turn divided into two regions with exactly the same size of population. To eliminate population shift effects, let us also assume a constant distribution of the population share in each of the two countries considered. In both A and B, the per capita income of one of the two regions is exactly twice that of the other region, and this situation remains unaltered for the whole of the period considered. There is, however, one major difference between these two countries. A is characterized by a high degree of regional mobility, such that, every year, its two regions switch their relative positions. The situation in B, however, differs in that the relative positions of its regions remain constant year on year. The type of analysis commonly found in the literature is essentially static in its approach, since it is based on cross sectional information and therefore will reveal no appreciable difference between A and B. In fact, given that there is no change over time in the cross sectional structure of the per capita income distribution of either of the countries, any inequality index that satisfies the properties of symmetry and scale independence will give exactly the same value for A and B throughout the period considered[2].

This example highlights the need to supplement standard inequality studies with additional data relating to the mobility of the distribution under analysis. It is precisely this issue that the present paper aims to address. Our objective is to analyze regional mobility in the spatial distribution of per capita income in the EU over the period 1977-1999. By this we hope to contribute to the understanding of the nature of observed territorial imbalances in the European context, and thereby draw some implications for EU regional policy makers. In fact, if a given level of inequality were found to be associated with a low degree of mobility, this would indicates that regions are becoming set in their relative positions. If so, this would reinforce the need for an active policy to reduce regional disparities, always supposing that the detected level of spatial inequality were found to be politically and socially inadmissible. If, however, the results of the analysis suggest that existing inequality can be largely explained by the variability of regional incomes, regional policy makers would need to focus their attention on ways to offset the adverse effects of economic cycles, and let traditional convergence policies take second place.

One of the main innovations in this paper lies in the instruments used to analyze regional mobility. Thus, most previous studies of this issue in the European setting are based on the distribution dynamics model proposed by Quah (1993; 1996a,b; 1997) (López-Bazo *et al.*, 1999; Cuadrado *et al.*, 2002; Le Gallo, 2004)[3]. However, in this paper we have added to the information provided by this methodology by calculating a set of measures used in the dynamic analysis of personal income distribution. Surprisingly, this approach has so far received little attention in the literature devoted to the analysis of territorial imbalances, due, in part, no doubt, to the limitations of the theoretical basis for the study of intra-distribution mobility. It is worth noting, however, that over the course of the last decade, major theoretical advances have been made in the analysis of intra-distribution mobility within this framework. In particular, there have been proposed a series of measuring procedures with similar axiomatic contents to those used in the study of inequality (Fields and Ok, 1999).

An analysis of the kind we wish to conduct requires that the data cover a representative sample of the economies within the area under study for a long enough time period. We have accomplished this by using the Cambridge Econometrics regional database, which has enabled us to employ statistical data on 197 NUTS2 regions for the period between 1977 and 1999[4].

The rest of the paper is structured as follows. Sections 2 and 3 examine the level and evolution of mobility in the regional distribution of per capita income in the EU using several complementary approaches. In order to complete these results, a non-parametric analysis based on the various instruments proposed by Quah (1996a,b; 1997) is performed in section 4. Subsequently, in section 5, we investigate the explanatory factors involved in regional mobility. Finally, the main conclusions of the paper are presented in section 6.

2 Mobility as compensation for inequality

We will begin our analysis of mobility by investigating its role in compensating for inequality. Traditionally, a high degree of mobility has been linked with lower long term inequality levels than tend to be detected in shorter sample periods. One way of testing mobility, therefore, is to observe the relationship between cross-sectional and longitudinal inequality.

Therefore, following common practice in the literature devoted to the dynamic analysis of personal income distribution, in this section we will consider the family of indices proposed by Shorrocks (1978a).

Let x_i^t be the per capita income of region i in period t, with $i = 1, 2, \ldots, n$, and $t = 1, 2, \ldots, T$. Accordingly, the associated per capita income distribution will be given by $x^t = (x_1^t, x_2^t, \ldots, x_n^t)$, while μ^t denotes the average per capita income in period t. Additionally, let $\hat{x}$ be the vector of aggregate per capita income over the T periods considered. That is, $\hat{x} = (\hat{x}_1, \hat{x}_2, ..., \hat{x}_n)$, where $\hat{x}_i = \sum\limits_{t=1}^{T} x_i^t$. Likewise, $\hat{\mu}$ stands for the average of $\hat{x}$.

We will now denote by $I(\hat{x})$ the set of inequality measures that are convex functions of the relative per capita incomes. Then, given the convexity of the function, it can be written as:

$$I(\hat{x}) = h\left(\frac{\hat{x}}{\hat{\mu}}\right) = h\left(\frac{\sum\limits_{t=1}^{T} x^t}{\hat{\mu}}\right) = h\left(\sum_{t=1}^{T} \omega^t \frac{x^t}{\mu^t}\right) \leq \sum_{t=1}^{T} \omega^t h\left(\frac{x^t}{\mu^t}\right) \tag{1}$$

where $\omega^t = \frac{\mu^t}{\hat{\mu}}$. Thus, from expression (1), it follows that:

$$I(\hat{x}) \leq \sum_{t=1}^{T} \omega^t I(x^t) \tag{2}$$

That is, the inequality index of the per capita incomes accumulated over the T periods considered can not exceed the weighted sum of the inequality indices of the individual periods. The rigidity index proposed by Shorrocks (1978a) is therefore defined as:

$$R(\hat{x}, x^t) = \frac{I(\hat{x})}{\sum\limits_{t=1}^{T} \omega^t I(x^t)} \tag{3}$$

with $R(\hat{x}, x^t) \leq 1$. Note that the above expression is valid only for inequality measures that are convex functions of the relative per capita incomes. This constraint does not impose a major drawback, however. Indeed, most of the indices commonly used in the literature (the Gini index, the family of Theil indices, Atkinson's indices, etc.) satisfy this property[5].

The index $R(\hat{x}, x^t)$ gives the value at which inequality diminishes as the time period considered is extended. Thus, for example, if $R(\hat{x}, x^t) = 0.90$, inequality over a given period will be 90 per cent of the average inequality corresponding to the set of subperiods contemplated. In other words, this index measures the stability of inequality as the sample period is increased. Indeed, if inequality remains stable as the period of reference is extended, we will have that:

$$I(\hat{x}) = \sum_{t=1}^{T} \omega^t I(x^t) \tag{4}$$

where $\frac{x^t}{\mu^t}$ is independent of t, such that $R(\hat{x}, x^t) = 1$. In other words, relative per capita incomes will not vary at all over time, which, in the present context, would indicate a lack of mobility. As long as there were some degree of intra-distribution mobility, however, more frequent and wider variations in relative incomes could be expected; that is, the value of $R(\hat{x}, x^t)$ would be less than one. Thus, $R(\hat{x}, x^t) = 0$ would indicate a case of perfect mobility in which $I(\hat{x}) = 0$[6]. Therefore, $R(\hat{x}, x^t)$ gives us the following measure of mobility:

$$RM(\hat{x}, x^t) = 1 - \frac{I(\hat{x})}{\sum_{t=1}^{T} \omega^t I(x^t)} \tag{5}$$

Figure 1 shows the results of the calculation of $RM(\hat{x}, x^t)$ for the EU regional distribution of per capita income between 1977 and 1999, taking different time periods ($m = 1, 2, \ldots, 23$). To check the sensitivity of the results to the choice of inequality index used to calculate $RM(\hat{x}, x^t)$, we have opted to incorporate into the analysis various measures of inequality, since each index features a different way of aggregating the information contained in the distribution (Sen, 1973). Following this approach, we selected the following measures: the Gini index, G, the family of Theil measures, $T(\beta)$ with $\beta = 0$ and $\beta = 1$, and the normative Atkinson index for different levels of inequality aversion, $A(\varepsilon)$ with $\varepsilon = 0.5$ and $\varepsilon = 2$ (see Chakravarty (1990) or Cowell (1995) for further details about these measures).

[INSERT FIGURE 1 AROUND HERE]

The results obtained show values of the mobility measure based on Shorrocks' rigidity index (1978a) that increase gradually as the period of reference is extended, irrespective of the inequality measure that is used (note that the ordinate axis has a scale of 0 to 0.1). This reveals that regional inequality in Europe declines very slowly when longer time intervals are considered. Thus, the influence of transient variability in regional disparities within the EU appears to be quite limited, therefore most of the observed inequality in this respect can be considered permanent. To illustrate this, Figure A1 displays the $R(\hat{x}, x^t)$ index for the whole of the 1977-1999 period. According to this, regional inequality in per capita income in the European context over the twenty-three years considered falls within a range of 93 to 99 per cent of average inequality for the set of subperiods contemplated, depending on the inequality index used to calculate $R(\hat{x}, x^t)$. This suggests that regional per capita income distribution in the EU is quite rigid and, therefore, barely mobile.

Nevertheless, detailed analysis of the information supplied in Figure 1 enables us to observe that the results obtained differ slightly according to the inequality index used in the calculation of $RM(\hat{x}, x^t)$. Both Theil indices follow a similar trend, though there appears to be a slight reduction in mobility as β increases. It is worth recalling, in this respect, that the β parameter captures the sensitivity of $T(\beta)$ to transfers between individuals at different points in the distribution. Following Shorrocks (1980), it can be shown that, as β increases, $T(\beta)$ actually becomes more sensitive to transfers at the upper end of the distribution. Also, as might be expected from the above results, mobility becomes greater as the value of ε increases. In fact, as is known, the higher the value of the inequality aversion parameter, the greater the sensitivity of Atkinson's index to what happens at the lower end of the distribution. The empirical evidence presented so far, therefore, appears to suggest that the reduction in inequality that takes place as the time interval is extended is greater in those European regions with lower per capita income levels.

3 Regional mobility: An analysis based on transition matrices

The measure of mobility considered in the previous section may in certain circumstances present some drawbacks relating to the significance of changes in the relative positions of the regions according to per capita income. To illustrate this problem, let us consider another example that highlights the multidimensional nature of mobility. Let us imagine a country with two regions, one of which enjoys some comparative advantage over the other, in terms, say, of its spatial location. In a situation of this kind, the region in question will, *ceteris paribus*, systematically register higher growth rates, giving rise to an increase in regional disparities, even after an initial situation of hypothetical equality. In other words, the rank ordering of the two regions will remain unaltered over time. In a context such as this, $RM(\hat{x}, x^t)$ will present positive values, though it could be argued that there is no mobility in the regional income distribution.

Keeping this fact in mind, in this section we use a new approach to the analysis of intra-distribution mobility, based on the observation of changes experienced by relative positions of the various regions.

One of the most intuitive options when approaching mobility studies in this way is to construct transition matrices. In order to define the concept of transition matrix, let us now suppose that the different regions in the distribution have been classified into m exhaustive and mutually exclusive classes according to their per capita income level. Further, let us imagine that we have data on the distribution of interest for two points in time, t_0 and t_1. In a case such as this, the matrix that summarizes the probabilities of regions shifting from one class to another between t_0 and t_1 is known as a transition matrix. Supposing, therefore, that the probabilities can be reasonably estimated from the corresponding relative frequencies, the transition matrix that shows changes in the distribution between t_0 and t_1 ($x^{t_0} \longrightarrow x^{t_1}$), will be the square matrix $\Pi(x^{t_0}, x^{t_1}) = [\pi_{jk}(x^{t_0}, x^{t_1})] \in \mathbb{R}_+^{m \times m}$, where

$\pi_{jk}(x^{t_0}, x^{t_1})$ denotes the proportion of regions that belonged to class j at t_0 and have shifted to class k at t_1. According to this definition, we have that $\sum_{k=1}^{m} \pi_{jk}(x^{t_0}, x^{t_1}) = 1$ for any $j = 1, 2, \ldots, m$. Therefore $\Pi(x^{t_0}, x^{t_1})$ is a stochastic matrix.

Numerous mobility measures based on transition matrices have been designed in the literature devoted to the dynamic study of personal income distribution (Geweke *et al.*, 1986). From this wide range of options we began by considering the following index based on Shorrocks (1978b)[7]:

$$SM^*(\Pi, \rho) = \frac{1 - \sum_{j=1}^{m} \rho_j \pi_{jj}}{1 - \frac{1}{m}} \tag{6}$$

where ρ_j denotes the population share of class j. This measure captures those aspects of the mobility concept that refer to the independence with regard to the initial situation. Nevertheless, $SM^*(\Pi, \rho)$ is of limited validity if the aim is to examine the dimension of mobility related to movement *per se* (Fields and Ok, 1999), since it is calculated exclusively from those elements that form the main diagonal of the transition matrix, thereby ignoring the rest of the elements in Π. To overcome this problem associated with the use of $SM^*(\Pi, \rho)$, we decided to consider a further index, proposed by Bartholomew (1973) and shown below:

$$BM^*(\Pi, \rho) = \sum_{j=1}^{m} \sum_{k=1}^{m} \rho_j \pi_{jk} |j - k| \tag{7}$$

The next step is to select an appropriate definition for each of the various classes. Faced with this problem, we decided to adopt a solution that enables us to obtain reasonably accurate information on regional movements across a sufficiently large number of groups, without risking any loss of representativity in the results. Thus, following the classification adopted by Cuadrado *et al.* (2002), we divided the regions that make up the distribution under analysis into five exhaustive and mutually exclusive classes. Each class was defined according to the regional per capita income in relation to the European average, which was assigned a value of 100: [0,75), [75,90), [90,110), [110,125) and [125,+∞).

[INSERT FIGURE 2 AROUND HERE]

Figure 2 shows the calculations of $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ after estimating the corresponding transition matrices. In addition, in order to isolate the effect of transient per capita income fluctuations associated with annual changes, we decided to use time periods of different length, thus we were also able to distinguish between short and medium term mobility.

The results, as expected, reveal that the longer the period of analysis the more mobility is observed in the regional per capita income distribution. Thus, on average, 91 per cent of the regions considered continued in the same class after a year. Taking the period as a whole, however, the percentage drops to 63 per cent. It is also worth stressing that the two mobility indices considered follow very similar trends. Given that the main difference between them lies in the different valuation given to shifts between classes, this result suggests a relatively low degree of intra-distribution mobility. Further confirmation of this is to be found in the various transition matrices estimated, which exhibit the highest values around the main diagonal[8]. This conclusion is consistent with the empirical evidence provided by Neven and Gouyette (1995) and López-Bazo *et al.* (1999) for a smaller geographical area and a shorter sample period than that considered in our paper.

Additionally, whatever index is used in the analysis, the results reveal a reduction in the mobility of the EU regional per capita income distribution between 1977 and 1999. Nevertheless, since mobility did not fall at an even rate over time, it is possible to identify a series of separate stages, each with its distinguishing features. Thus, the main reduction in $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ took place between 1977 and the early eighties. From then onwards, however, there was a change of trend leading to an increase in regional mobility that continued until the end of that decade. During the early nineties, there was a further decrease that was followed by a new stage characterized by a rise in $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$[9].

In this context, however, it is necessary to stress that the above results cannot be valued normatively without taking into account the degree of inequality observed in the distribution under analysis. In this respect, numerous studies have coincided in reporting a lack of regional convergence in per capita income in the European context from the mid-seventies onwards (Neven and Gouyette, 1995; López-Bazo *et al.*, 1999; Rodríguez-Pose, 1999). The analysis performed in this section, meanwhile, shows that this persistence in regional disparities has coincided in time with a process of consolidation in the relative positions of the various regions, which stresses the need for an active regional policy at European level, assuming that the existing level of spatial inequality is considered politically and socially unacceptable[10].

Finally, in light of the volatility of $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ in short-term observations, we explored the relationship between the economic cycle and regional mobility trends in the European context. To this end we estimated the statistical correlation between per capita income growth rates in the EU and annual fluctuations in the two mobility measures considered in this section (Pekkala, 2000). This exercise was then repeated incorporating the assumption that the economic cycle has a delayed effect on regional mobility (Fischer and Nijkamp, 1987). In both cases, however, the correlation coefficients, though positive, were not statistically significant.

4 A non-parametric analysis of intra-distribution mobility

By means of the various tools employed in the preceding section, we have explored the level and evolution of regional mobility in the EU between 1977 and 1999. It is necessary to bear in mind, however, that $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ were calculated on the basis of the information supplied by various transition matrices, obtained by dividing the distribution of interest into a series of exhaustive and mutually exclusive classes. However, since there is no procedure for finding the optimal number of classes in each case, the researcher is obliged to make an arbitrary decision in this respect (Quah, 1993, 1996a; Kremer *et al.*, 2001).

To address this problem, Quah (1996a, 1997) suggests substituting the transition matrix with a stochastic kernel that reflects the probabilities of transition between a hypothetically infinite number of classes (Durlauf and Quah, 1999). The stochastic kernel can be reached by estimating the density function of the distribution over a given period, $t+k$, conditioned by the values of a previous period, t. Specifically, the joint density function of the distribution at t and $t+k$ is estimated non-parametrically and normalized by the implicit marginal distribution at t in order to obtain the corresponding conditional probabilities.

Figure 3 shows the stochastic kernel estimated for the European regional per capita income distribution over a period of twenty-three years ($t = 1977$ and $t+k = 1999$). Gaussian kernel functions were used, while the smoothing parameter values were selected following Silverman (1986, p. 86). The three-dimensional graph informs about the probabilities associated with each pair of values in the first and last years of the study period. In other words, the stochastic kernel provides, in a way analogous to that of a discrete transition matrix, the probability distribution of 1999 per capita income for regions with a given per capita income in 1977. Thus, if the probability mass is concentrated around the main diagonal, the intra-distribution dynamics are characterized by a high level of persistence in the relative positions of the regions over time and, therefore, low mobility. If, on the other hand, the density is located mainly on the opposite diagonal to the main diagonal, this would indicate that regions at each end of the distribution exchange their relative positions throughout the period. Finally, the probability mass could, in theory, accumulate parallel to the t axis. This would reflect the existence of a process of convergence around a given level of per capita income. In order to aid interpretation of the graph, Figure 3 also includes a contour plot on which the lines connect points at the same height on the three-dimensional kernel.

[INSERT FIGURE 3 AROUND HERE]

The results obtained fully uphold the conclusions reached in the previous analysis based on the data from the discrete transition matrices. Indeed, as can be seen from Figure 3, the mass of probability is concentrated around the main diagonal. As we are already aware, this shows that there was little mobility in the distribution of regional per capita income between

1977 and 1999. There is a general tendency, therefore, for the European regions to maintain their relative positions throughout the twenty-three years contemplated. Nevertheless, the analysis carried out shows that the mobility patterns of the various regions vary according to their economic development level. It is possible to observe, for example, how regions with a per capita income close to the European average exhibit a relatively higher degree of mobility over time, while those located at both ends of the distribution are characterized by a stronger persistence in their relative positions. In particular, the information provided by Figure 3 in this respect confirms that there is comparatively less mobility among more highly developed regions than among low per capita income regions over the time period considered[11].

In light of these results, we completed the above analysis with further information relating to the behavior of the regions situated at the two ends of the distribution under consideration, taking these to be the ones in which per capita income fell outside the interval of 50 to 150 per cent of the European average. Our calculations indicate that 27 per cent of the regions with a per capita income below 50 per cent of the European average in 1977, remained in the same situation in 1999. In fact, of the 22 regions whose per capita income in 1977 was below 50 per cent of the European average, only the Portuguese regions of Norte, Centro, Alentejo, Algarve, Açores and Madeira remained in the same situation twenty-three years later. However, out of the other 16, only the Spanish regions of Aragón, Baleares, Madrid, Cataluña and La Rioja had succeeded in raising their per capita income above 75 per cent of the European average, which is further support for the results obtained earlier. There is a different situation at the upper end of the distribution, however, where, out of the 13 regions who began the period with a per capita income above 150 per cent of the European average, only the Swedish regions of Norra Mellansverige, Mellersta Norrland and Ovre Norrland, together with Valle d'Aosta and Groningen had fallen below that level by 1999, though none of them dropped below 125 per cent of the European average.

5 Explanatory factors for regional mobility

To round off the results obtained in the previous sections, we next investigated the role played by a series of factors in accounting for the observed level of intra-distribution mobility in the EU from 1977 to 1999. Our specific aim was to ascertain why the relative position of some regions improved, while that of others deteriorated over the twenty-three years considered.

Thus, our first step was to determine which dependent variable to use in the analysis. In this respect it is worth noting that, for the study period considered, the use of data deriving from one of the various mobility measures calculated in the preceding pages would leave us, at best, with only twenty-two values for each index. Needless to say, even if we were willing to consider only interannual mobility, such a degree of freedom would be clearly insufficient. To address the problems surrounding this issue, we opted for the alternative of considering an individual measure of relative regional mobility, MOB_i, based on the variation over time in the per capita income of each region normalized according to the European average. Unlike standard mobility indices, MOB_i informs about the direction of regional shifts, so that it is possible to determine which regions have improved or fallen back in their relative position over the sample period. Additionally, it is worth mentioning that MOB_i has the advantage of allowing us to capture the distribution dynamics that leave the regional ranking unaltered.

Having established the dependent variable, we investigated how far the initial levels of per capita income ($GVApc_{i0}$), per capita expenditure in investment ($INVpc_{i0}$), and population density (DEN_{i0}), contribute to explain the observed regional mobility pattern in the European setting. We also explored the role played by economic geography in this context. For this, we calculated a market potential index for each region at the beginning of the sample period, in order to control for the impact of market access on distribution dynamics. In this way, we aimed to take into account the fact that the potential demand for goods and services in a given location is influenced by its accessibility to consumers (Harris, 1954). The new economic geography models lend theoretical support for the use of an index of this type (Krugman, 1992), while various recent studies have underlined its empirical relevance (Maurseth, 2001). In this paper we defined the market potential of each region as the inverse-distance weighted sum of the purchasing power of all other regions, in order to capture the effect of transport costs (MP_{i0}).

Moreover, the importance of the role of the sectoral composition of economic activity in regional growth processes is widely known (for the European case, see Paci and Pigliaru (2000) or Gil *et al.* (2002)). It is therefore reasonable to assume that the initial productive structure and its subsequent evolution may be related to modifications in the relative positions of the regions. Taking this idea into account, we decided to introduce into our model the initial share in regional employment of agriculture (EAG_{i0}), the financial sector (EFS_{i0}), and non-market services ($ENMS_{i0}$); together with the variation in these variables over the period analyzed (ΔEAG_i, ΔEFS_i and $\Delta ENMS_i$). It is in fact common practice in the literature devoted to the estimation of convergence equations to include a variable to capture the size of the agricultural sector, in order to control for differences in the sectoral composition of activity across the different territorial units to be analyzed. However, bearing in mind the process of increasing tertiarization that has been taking place in the European economy over the last few decades (European Commission, 1999), we also took into account the role played in this context by advanced services and public employment, which we approximated, respectively, with EFS_{i0}, ΔEFS_i, $ENMS_{i0}$ and $\Delta ENMS_i$.

Thus, our proposed model to explain the regional mobility pattern in the EU between 1977 and 1999 is defined as follows:

$$\begin{aligned} MOB_i &= \beta_0 + \beta_1 GVApc_{i0} + \beta_2 INVpc_{i0} + \beta_3 DEN_{i0} + \beta_4 MP_{i0} + \beta_5 EAG_{i0} \\ &\quad + \beta_6 \Delta EAG_i + \beta_7 EFS_{i0} + \beta_8 \Delta EFS_i + \beta_9 ENMS_{i0} + \beta_{10} \Delta ENMS_i \\ &\quad + u_i \qquad (8) \end{aligned}$$

where u_i is a random disturbance.

Table 1 shows the estimation of the above model by ordinary least squares (OLS) for different time periods. Before interpreting the results, however, it should be borne in mind that several studies have underlined the relevance of the spatial dimension in explaining observed territorial imbalances in per capita income in the EU (López-Bazo *et al.*, 1999; Le Gallo and Ertur, 2003; Maza and Villaverde, 2004). The results of the cited works clearly show that the development level is not randomly distributed across the European territory. On the contrary, the available empirical evidence reveals the presence of positive spatial dependence in this context, which suggests that neighboring regions tend to be characterized by similar per capita income levels.

[INSERT TABLE 1 AROUND HERE]

In order to assess the importance of this issue within the context of our paper, we defined a spatial weight matrix, W, to capture the degree of interdependence between each pair of regions i and j. A first option is to use the concept of first order contiguity, according to which $w_{ij} = 1$ if regions i and j are physically adjacent and 0 otherwise. This is in fact the option taken, among others, by López-Bazo *et al.* (1999) or Rey and Montouri (1999). However, the use of this type of matrix may raise problems in the European context, where the presence of islands means that W will include rows and columns containing only zeros. This means omitting the observations in question from the analysis, which in turn has an effect on the interpretation of the results obtained. In this paper, therefore, we opted for the alternative of using a spatial weight matrix that takes into account interactions taking place beyond the adjacent regions. Specifically, following the proposal made by Anselin and Bera (1998), we considered a row-standardized spatial weight matrix based on the square inverse distance between the centroids of the various regions.

Next, various spatial autocorrelation tests, namely, the Moran's I (Cliff and Ord, 1972), the Lagrange multiplier tests for the spatial error and the spatial lag models proposed respectively by Burridge (1980) and Anselin (1988a), plus their robust versions (Anselin *et al.*, 1996), were computed from the OLS residuals. The results of these tests led to the rejection of the null hypothesis of absence of residual spatial dependence, suggesting, in accordance with Florax and Folmer (1992), the need to include the spatial lag of the dependent variable, $WMOB_i$, in the list of regressors. That is,

$$\begin{aligned} MOB_i &= \beta_0 + \beta_1 GVApc_{i0} + \beta_2 INVpc_{i0} + \beta_3 DEN_{i0} + \beta_4 MP_{i0} + \beta_5 EAG_{i0} \\ &\quad + \beta_6 \Delta EAG_i + \beta_7 EFS_{i0} + \beta_8 \Delta EFS_i + \beta_9 ENMS_{i0} + \beta_{10} \Delta ENMS_i \\ &\quad + \gamma WMOB_i + u_i \end{aligned} \tag{9}$$

where γ is the spatial autoregressive parameter (spatial lag model).

Nevertheless, the estimation of model (9) by OLS is inconsistent, due to simultaneity induced by the spatial lag. Instrumental variables and maximum likelihood estimators have been suggested to provide consistent estimates (Anselin, 1988b). Taking this into account, the maximum likelihood estimates (ML) of the spatial lag model (9) for the different time intervals considered are reported in the second, fourth and sixth columns in Table 1.

Table 1 shows that $WMOB_i$ is significant and positive for the 1977-1999 period. In fact, this result is confirmed if we take into account the information provided by the various tests on the spatial autoregressive parameter carried out. All of this clearly shows that the behavior of neighboring regions has a positive effect in explaining the variability of the dependent variable, which is coherent with the results obtained by Le Gallo (2004).

Likewise, our estimates reveal the presence of a negative relationship between MOB_i and the initial per capita income level, which allows us to complete and qualify some of the findings from the analysis performed in the preceding section. It is also worth noting the low dynamism of the agricultural regions. Indeed, the presence in 1977 of a relatively large agricultural sector or the growth of this activity in employment terms, are found to be associated with a worsening of the relative positions of the various regions. Meanwhile, EFS_{i0} is also statistically significant. This suggests that upward regional shifts are positively linked to the employment share of certain types of advanced high-productivity services. In any event, the increase in non-market services is negatively correlated with MOB_i. This result is consistent with the empirical evidence presented by Rodríguez-Pose and Fratesi (2004a), who stress the fact that the European peripheral regions, characterized by high levels of public employment, presented more moderate growth rates than the rest between 1980 and 2000. Finally, it is worth mentioning that $INVpc_{i0}$, DEN_{i0} and MP_{i0} are not statistically significant. This suggests that per capita expenditure in investment, population density and market potential do not contribute to explain the distribution dynamics.

Next, in order to detect possible variations in performance over time, we decided to repeat the analysis for a number of shorter intervals. The results for the 1977-1988 subperiod are very similar to those just discussed for the period as a whole, however. In this particular case, the only difference arises from the fact that the increase in employment in advanced services appears to have a negative effect on the dependent variable.

When analyzing the 1988-1999 subperiod, we introduced a slight modification to the model we had been estimating so far, in order to obtain a first impression of the relationship between the EU regional policy and observed intra-distribution mobility. This involved the inclusion of a dummy variable, $RO1_i$, to enable us to identify all the regions that held Objective 1 status in any of the various programming periods. In this way we expected to see whether regions that had benefited from priority treatment under EU regional policy after the Structural Funds reform in 1988 performed differently from the rest. In this respect, the information provided by Table 1 suggests that $RO1_i$ is unrelated to variations in the dependent variable. This result should nevertheless be interpreted with caution. On the one hand, it should be borne in mind that both Boldrin and Canova (2001) and Rodríguez-Pose and Fratesi (2004b) insist on the low mobility of the less developed regions of the EU during the nineties. Nevertheless, there are some exceptions to this general trend that should not be overlooked. This is the case of Southern and Eastern Ireland or the Abruzzi in Italy, for example. Accordingly, it would be extremely risky to judge something as complex as the relationship between EU regional policy and the dynamics exhibited over the course of the last decade by the Objective 1 regions exclusively on the basis of the results of an analysis of this nature.

Finally, with respect to the rest of the explanatory variables considered in this study, the main difference between the estimations for the 1988-1999 subperiod and those for the period as a whole is the lack of statistical significance of $GVApc_{i0}$ and EFS_{i0} during those 12 years. In other words, it is not possible to establish any link between the changes that have taken place in the relative situation of the various regions in development terms and their initial per capita income level or share of the financial sector in total employment in 1988. At the same time, however, it is worth noting that increases in employment in that sector over this twelve-year period had a positive impact on the dependent variable.

Next, we examined the role played by national borders in explaining the spatial autocorrelation detected previously in this context. The above analysis was therefore repeated for an alternative definition of the spatial weight matrix used to capture interdependence between the various regions contemplated. In particular, the spatial weight matrix used so far was modified, so that all weights corresponding to regions of different countries are set equal to zero[12].

[INSERT TABLE 2 AROUND HERE]

The results of this exercise are shown in Table 2. Comparison with the information displayed in Table 1 reveals that the statistically significant values obtained in the spatial autocorrelation tests are in all cases higher in Table 2. This highlights the importance of the national component in explaining the regional mobility observed in the European Union, which is consistent with the conclusions of various studies on existing regional disparities in the European context (Quah, 1996c; Ezcurra *et al.*, 2005). Notwithstanding this circumstance, the estimations are, in any event, very similar to those discussed earlier. In fact, the only noteworthy discrepancies are found in the 1988-1999 sub-period. Thus, in contrast to the situation described by Table 1, the variation in employment share in advanced services has no impact on the dependent variable, while per capita expenditure in investment and

MOB_i are negatively correlated.

6 Conclusions

The regional mobility in the spatial distribution of per capita income in the EU over the period 1977-1999 has been examined in this paper using a number of complementary approaches. We began by calculating a set of measures used in the literature devoted to the dynamic analysis of personal income distribution. Our results show a decline in mobility within the distribution of interest over the period of observation. A further feature of note is the relatively low level of intra-distribution mobility. This conclusion is in fact confirmed when stochastic kernels and contour plots are estimated for a number of intervals of different lengths. Thus, with only a few exceptions, the European regions tended to maintain their relative positions over the twenty-three years considered.

Our results also show that regional mobility patterns vary as a function of economic development. In fact, the regions with a per capita income close to the European average tended to register a relatively higher mobility degree over time, while those at either end of the distribution were characterized by a stronger persistence in their relative positions. In particular, the less developed regions showed greater mobility than those located at the upper end of the distribution.

Finally, we carried out a regression analysis in order to identify the explanatory factors of regional mobility in the EU. The results obtained for the 1977-1999 period reveal a negative relationship between upward mobility and initial per capita income levels. Furthermore, the presence at the beginning of the period of a relatively large agricultural sector or the increase in the share of employment in this sector are found to be associated with a falling back of the relative position of the region in question. In fact, an increase in employment in non-market services has a similar effect, in contrast to what happens with the initial importance of advanced high-productivity services. Likewise, the Objective 1 regions failed in general terms to improve their relative situation over the 1988-1999 period, in spite of the priority treatment they received under EU regional policy. In any event, it is worth noting that our estimates highlight the major role played by the mobility performance of neighboring regions in explaining the distribution dynamics.

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Notes

[1]Article 2 of the Treaty of the EU specifically states that "The Community shall have as its task (...), to promote (...) a harmonious, balanced and sustainable development of economic activities, (...) sustainable and non-inflationary growth, (...) a high degree of competitiveness and convergence of economic performance (...)".

[2]The properties of symmetry and scale independence do not constitute a major limitation. Indeed both are basic properties that any inequality index can reasonably be expected to fulfill (Cowell, 1995). In any event, for the purposes of our example, we can overcome the need for the inequality index to satisfy the property of scale independence by simply assuming the average per capita incomes of A and B to be equal.

[3]See Durlauf and Quah (1999) for further details on the theoretical basis for the dynamic analysis of spatial distributions.

[4]Lack of complete series, however, has obliged us to eliminate from the analysis the Member States incorporated into the EU in May 2004, the *Länder* of former East Germany, The French overseas departments and the Spanish territories in North Africa.

[5]The most outstanding exception is the variance of the logarithms.

[6]If $I(\hat{x}) = 0$, it follows that $\hat{x}_1 = \hat{x}_2 = \ldots = \hat{x}_n$.

[7]The mobility measure proposed by Shorrocks (1978b) is given by:

$$SM(\Pi) = \frac{m - tr(\Pi)}{m - 1}$$

where $tr(\Pi)$ denotes the trace of the matrix Π. Note that, in contrast to what occurs with $SM^*(\Pi, \rho)$, this index assigns identical weight to each of the m classes. Indeed, if $\rho_j = \frac{1}{m}$ for any $j = 1, 2, \ldots m$, then $SM^*(\Pi) = SM(\Pi)$.

[8]The medium and full term transition matrices are included in the appendix. The rest, which are not shown for lack of space, are available from the authors upon request.

[9] In order to test the robustness of the above results, we recalculated $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$ for an eight-category classification of the European regions, based on the following per capita income levels: [0,50), [50,75), [75,90), [90,100), [100,110), [110,125), [125,150) and [150,+∞). The results, which are shown in the appendix, are very similar to those just discussed.

[10] Note that, for a given level of inequality, high mobility would be a sign of strong cyclical variability in regional incomes. In this kind of context, regional policy should switch the focus away from the objectives of traditional convergence policies and direct it towards the need to mitigate the adverse effects of economic cycles.

[11] In order to test the robustness of the results, we decided to repeat the above analysis using data only for the subperiods 1977-1988 and 1988-1999. The results, shown in the appendix, are very similar to those discussed in this section.

[12] We had to exclude Denmark and Luxembourg from this analysis, given that these two countries are formed by a single NUTS2 region according to the Eurostat territorial classification.

Appendix

[INSERT FIGURE A1 HERE]

[INSERT FIGURE A2 HERE]

[INSERT FIGURE A3 HERE]

[INSERT FIGURE A4 HERE]

[INSERT TABLE A1 HERE]

[INSERT TABLE A2 HERE]

[INSERT TABLE A3 HERE]

List of Figures and Tables

Figure 1: Regional mobility measured by $RM(\hat{x}, x^t)$ (for various inequality measures).

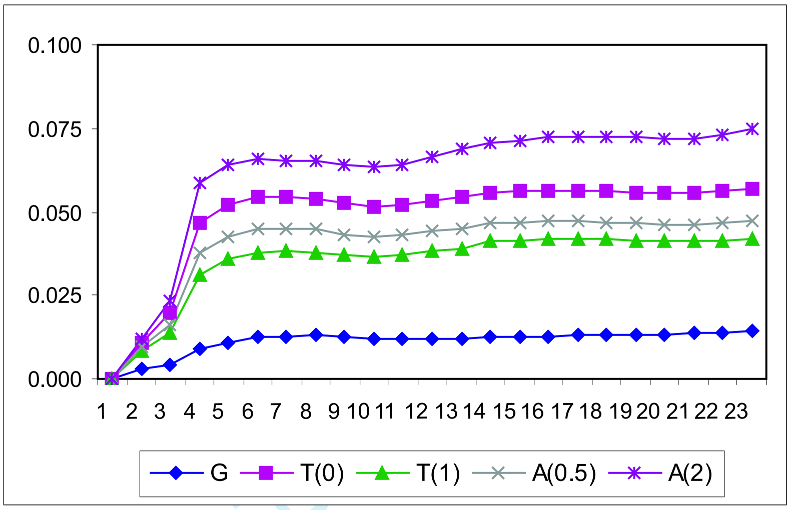

Figure 2: Regional mobility measured by $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$, $m = 5$.

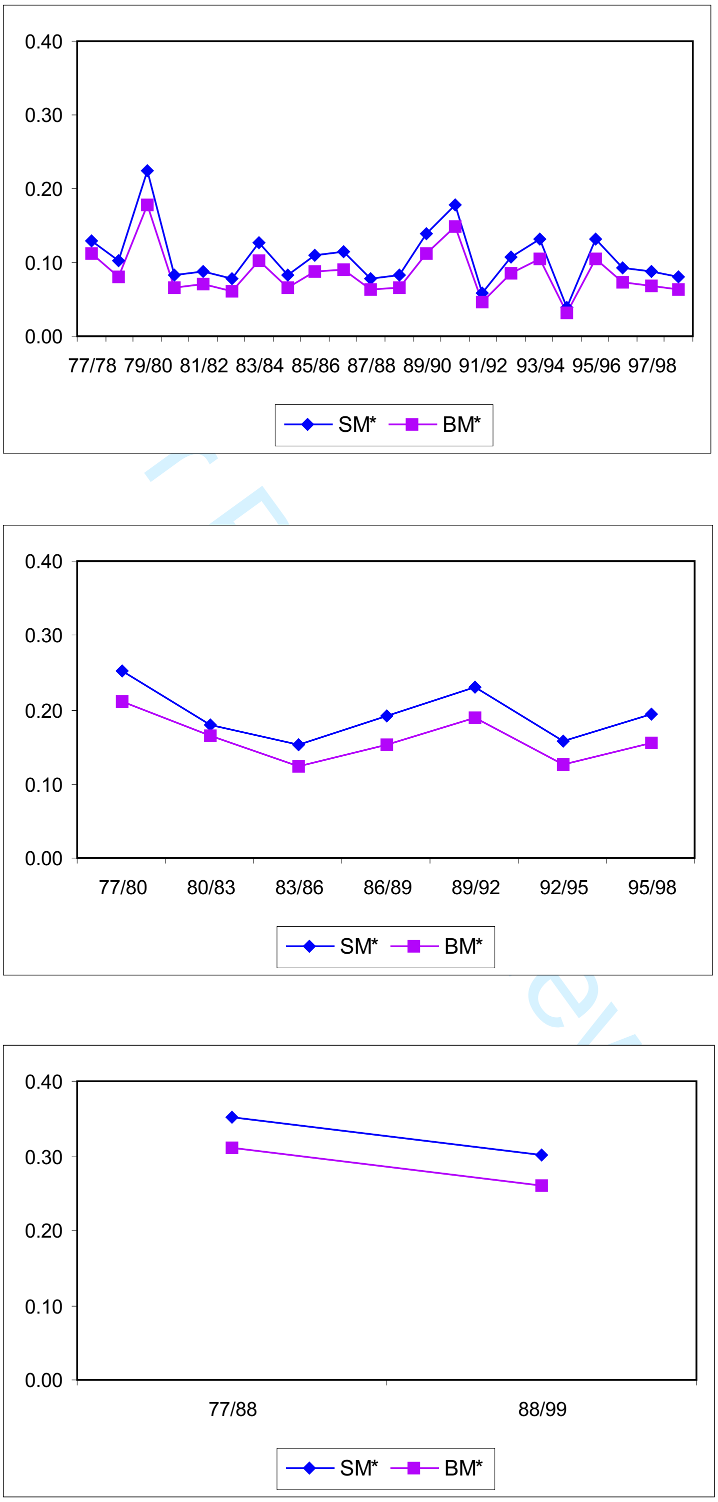

Figure 3: Stochastic Kernel and contour plot of the regional per capita income distribution, 1977-1999 (European average=100).

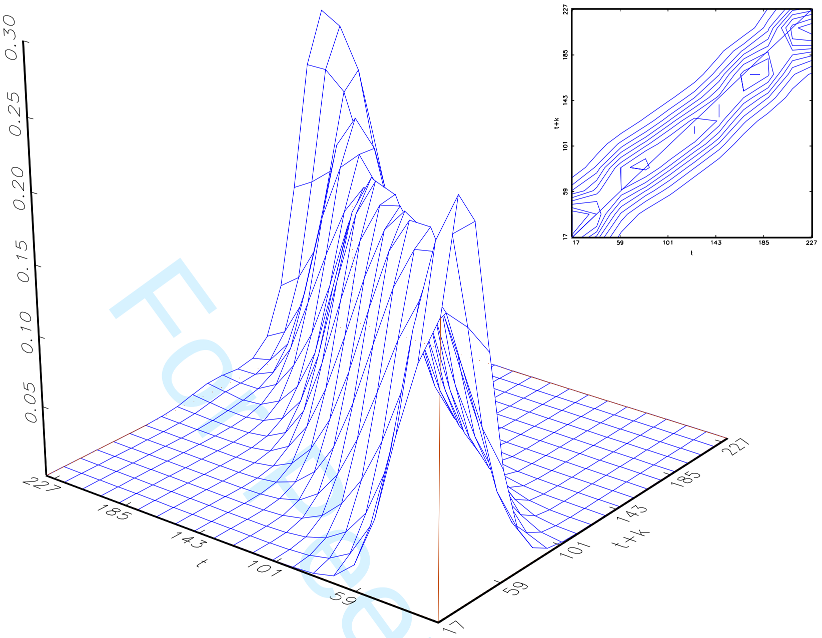

Figure A1: Shorrocks' rigidity index ($R(\hat{x}, x^t)$), 1977-1999.

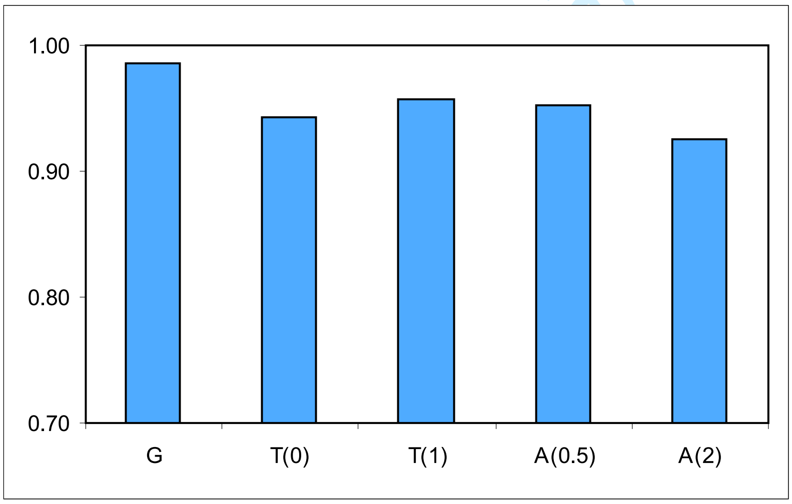

Figure A2: Regional mobility measured by $SM^*(\Pi, \rho)$ and $BM^*(\Pi, \rho)$, $m = 8$.

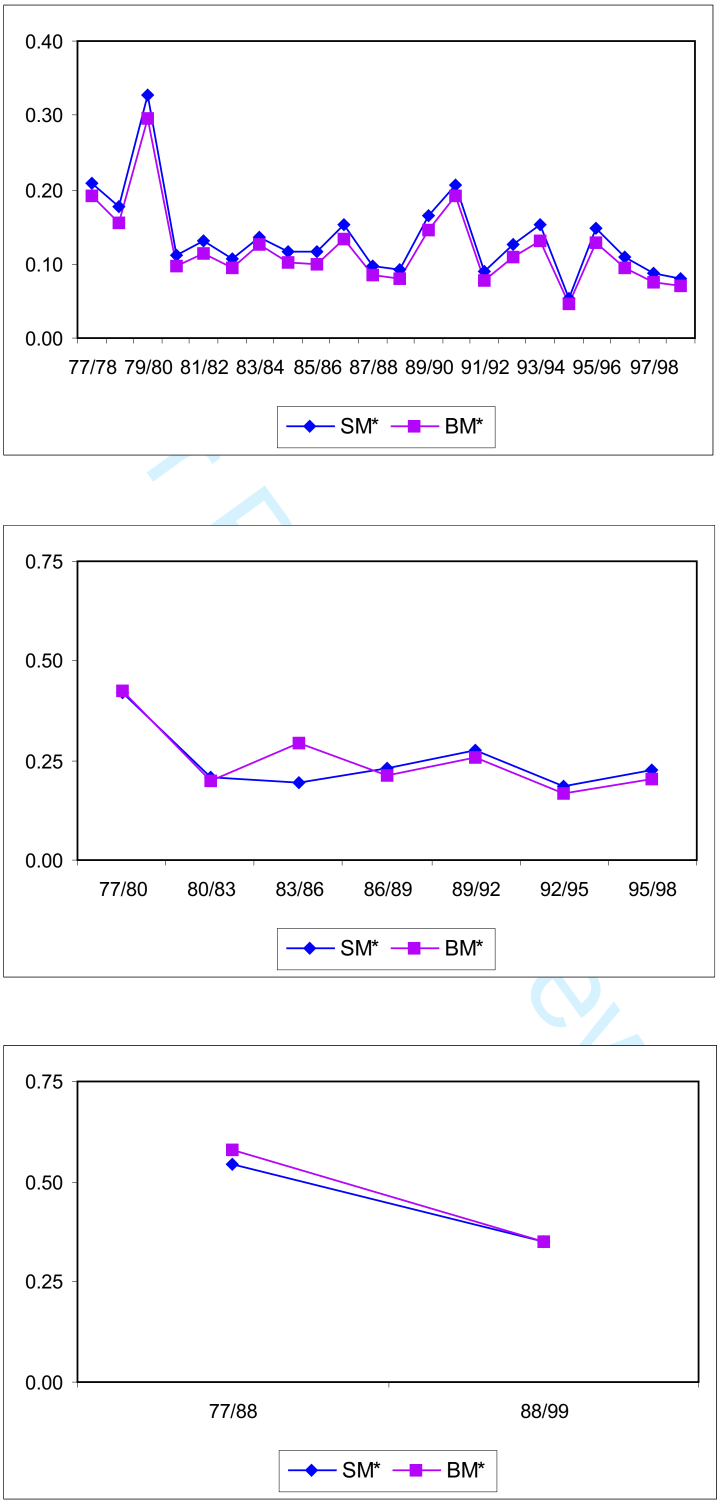

Figure A3: Stochastic kernel and contour plot of the regional per capita income distribution, 1977-1988 (European average=100).

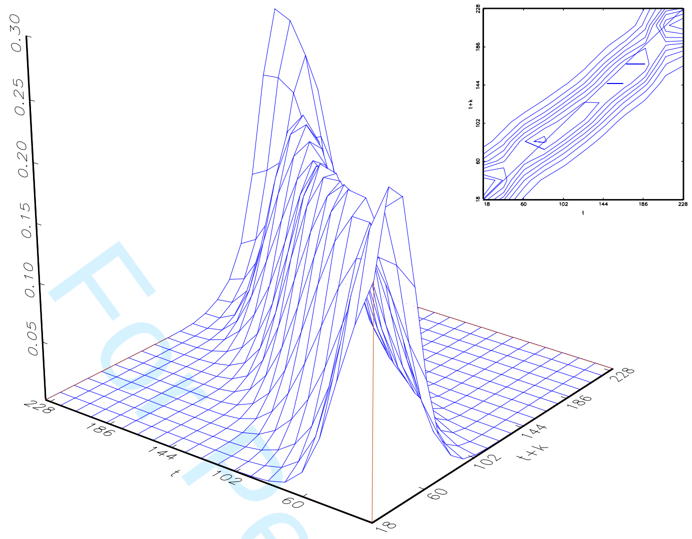

Figure A4: Stochastic kernel and contour plot of the regional per capita income distribution, 1988-1999 (European average=100).

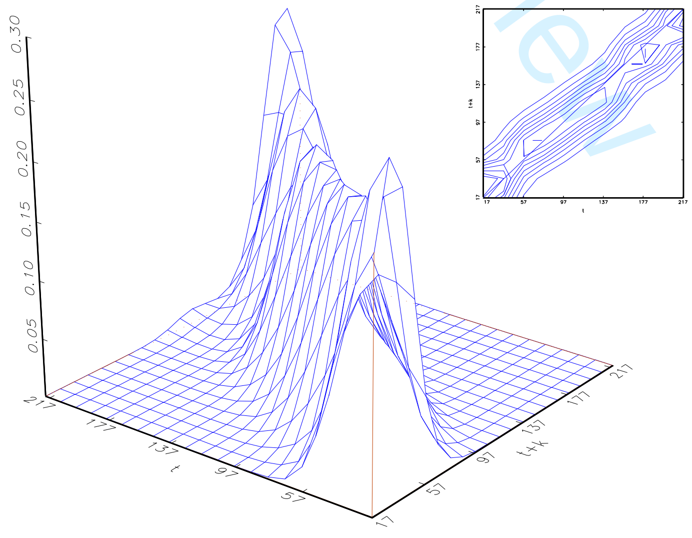

Table 1: Explanatory factors of mobility.

Dependent var.	MOB_i					
Period	1977-1999		1977-1988		1988-1999	
Independent var.	OLS	ML-LAG	OLS	ML-LAG	OLS	ML-LAG
$Constant$	0.281	0.239	0.318	0.226	0.035	0.054
	(0.138)	(0.059)	(0.036)	(0.067)	(0.609)	(0.336)
$GVApc_{i0}$	-0.032	-0.022	-0.031	-0.023	-0.001	-0.002
	(0.000)	(0.000)	(0.000)	(0.000)	(0.514)	(0.280)
$INVpc_{i0}$	-0.000	-0.002	0.000	0.000	-0.000	-0.001
	(0.950)	(0.734)	(0.928)	(0.994)	(0.922)	(0.682)
DEN_{i0}	-0.012	-0.013	-0.002	-0.003	-0.012	-0.011
	(0.256)	(0.126)	(0.762)	(0.598)	(0.078)	(0.146)
MP_{i0}	-0.051	-0.083	-0.080	-0.032	-0.043	-0.058
	(0.716)	(0.346)	(0.416)	(0.641)	(0.323)	(0.122)
EAG_{i0}	-0.019	-0.008	-0.010	-0.005	-0.005	-0.003
	(0.000)	(0.007)	(0.000)	(0.028)	(0.001)	(0.026)
ΔEAG_i	-0.031	-0.013	-0.026	-0.013	-0.010	-0.005
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.051)
EFS_{i0}	0.069	0.057	0.073	0.059	0.013	0.015
	(0.000)	(0.000)	(0.000)	(0.000)	(0.158)	(0.106)
ΔEFS_i	-0.044	-0.003	-0.069	-0.048	0.074	0.058
	(0.295)	(0.916)	(0.006)	(0.001)	(0.011)	(0.035)
$ENMS_{i0}$	0.002	0.000	0.001	0.001	0.000	-0.000
	(0.520)	(0.868)	(0.627)	(0.726)	(0.727)	(0.898)
$\Delta ENMS_i$	-0.010	-0.012	-0.017	-0.017	-0.008	-0.005
	(0.012)	(0.000)	(0.003)	(0.002)	(0.000)	(0.006)
RO_i					-0.006	-0.011
					(0.783)	(0.543)
$WMOB_i$		0.826		0.705		0.704
		(0.000)		(0.000)		(0.000)
LIK	77.101	115.337	126.983	157.797	215.638	230.909
AIC	-132.201	-204.673	-231.965	-289.594	-407.277	-433.819
SC	-96.085	-161.991	-195.851	246.912	-367.878	-387.855
Moran's I	10.166		8.241		7.365	
	(0.000)		(0.000)		(0.000)	
LMERR	65.717		40.779		32.027	
	(0.000)		(0.000)		(0.000)	
R-LMERR	2.339		0.411		0.069	
	(0.126)		(0.521)		(0.793)	
LMLAG	95.873		80.516		40.938	
	(0.000)		(0.000)		(0.000)	
R-LMLAG	32.495		40.148		8.980	
	(0.000)		(0.000)		(0.000)	
Wald test on γ		83.929		28.147		37.587
		(0.000)		(0.000)		(0.000)
LM test on γ		95.873		80.156		40.938
		(0.000)		(0.000)		(0.000)

Notes: W is based on the square inverse distance between the various regions. p-values are in parentheses. Robust standard errors were computed according to White (1980, 1982). LIK is the value of the log of the maximum likelihood function, while AIC and SC are respectively the Akaike and Schwarz information criteria. LMERR is the Lagrange multiplier test for residual spatial autocorrelation and R-LMERR is its robust version. LMLAG is the Lagrange multiplier test for the spatially lagged dependent variable and R-LMLAG is its robust version. The null hypothesis in the Wald and Lagrange multiplier (LM) tests on the spatial autoregressive parameter is that $\gamma = 0$.

Table 2: Explanatory factors of mobility and national borders.

Dependent var.	MOB_i					
Period	1977-1999		1977-1988		1988-1999	
Independent var.	OLS	ML-LAG	OLS	ML-LAG	OLS	ML-LAG
Constant	0.273	0.151	0.325	0.184	0.037	0.036
	(0.122)	(0.205)	(0.030)	(0.142)	(0.563)	(0.387)
$GVApc_{i0}$	-0.032	-0.020	-0.031	-0.020	-0.001	-0.009
	(0.000)	(0.000)	(0.000)	(0.000)	(0.561)	(0.613)
$INVpc_{i0}$	-0.002	-0.004	-0.000	-0.001	-0.001	-0.003
	(0.688)	(0.358)	(0.951)	(0.898)	(0.526)	(0.034)
DEN_{i0}	-0.007	-0.010	-0.001	-0.001	-0.010	-0.008
	(0.420)	(0.162)	(0.932)	(0.781)	(0.083)	(0.122)
MP_{i0}	-0.018	0.001	-0.007	0.003	-0.041	-0.022
	(0.896)	(0.995)	(0.432)	(0.967)	(0.333)	(0.478)
EAG_{i0}	-0.018	-0.007	-0.010	-0.005	-0.005	-0.002
	(0.000)	(0.011)	(0.000)	(0.032)	(0.001)	(0.036)
ΔEAG_i	-0.031	-0.013	-0.026	-0.013	-0.011	-0.004
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.053)
EFS_{i0}	0.062	0.045	0.072	0.049	0.007	0.005
	(0.000)	(0.000)	(0.000)	(0.000)	(0.449)	(0.459)
ΔEFS_i	-0.089	-0.024	-0.089	-0.039	0.057	0.026
	(0.036)	(0.387)	(0.005)	(0.026)	(0.055)	(0.278)
$ENMS_{i0}$	0.003	0.002	0.002	0.001	0.001	0.000
	(0.289)	(0.298)	(0.523)	(0.433)	(0.454)	(0.849)
$\Delta ENMS_i$	-0.010	-0.010	-0.017	-0.017	-0.008	-0.004
	(0.016)	(0.001)	(0.004)	(0.003)	(0.000)	(0.016)
RO_i					-0.009	-0.017
					(0.686)	(0.302)
$WMOB_i$		0.602		0.553		0.621
		(0.000)		(0.000)		(0.000)
LIK	81.185	125.947	125.567	166.601	219.075	254.089
AIC	-140.370	-225.893	-229.131	-307.203	-414.151	-480.179
SC	-104.367	-183.344	-193.128	-264.654	-374.875	-434.357
Moran's I	10.698		8.846		10.041	
	(0.000)		(0.000)		(0.000)	
LMERR	81.199		53.804		70.704	
	(0.000)		(0.000)		(0.000)	
R-LMERR	1.968		0.059		7.497	
	(0.161)		(0.808)		(0.006)	
LMLAG	115.172		104.202		105.427	
	(0.000)		(0.000)		(0.000)	
R-LMLAG	35.941		50.458		42.220	
	(0.000)		(0.000)		(0.000)	
Wald test on γ		72.610		37.987		95.257
		(0.000)		(0.000)		(0.000)
Wald test on γ		115.172		104.202		105.427
		(0.000)		(0.000)		(0.000)

Notes: W is based on the square inverse distance between the regions of the same country. p-values are in parentheses. Robust standard errors were computed according to White (1980, 1982). LIK is the value of the log of the maximum likelihood function, while AIC and SC are respectively the Akaike and Schwarz information criteria. LMERR is the Lagrange multiplier test for residual spatial autocorrelation and R-LMERR is its robust version. LMLAG is the Lagrange multiplier test for the spatially lagged dependent variable and R-LMLAG is its robust version. The null hypothesis in the Wald and Lagrange multiplier (LM) tests on the spatial autoregressive parameter is that $\gamma = 0$.

Table A1: Transition matrix, 1977-1988.

	Per capita income classes				
Number of regions	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
46	0.81	0.17	0.02	0.00	0.00
45	0.18	0.67	0.15	0.00	0.00
46	0.07	0.13	0.71	0.07	0.02
24	0.00	0.00	0.33	0.63	0.04
36	0.00	0.00	0.03	0.22	0.75

Table A2: Transition matrix, 1988-1999.

	Per capita income classes				
Number of regions	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
48	0.98	0.00	0.02	0.00	0.00
44	0.15	0.55	0.30	0.00	0.00
50	0.00	0.16	0.68	0.16	0.00
26	0.00	0.00	0.15	0.73	0.12
29	0.00	0.00	0.07	0.10	0.83

Table A3: Transition matrix, 1977-1999.

	Per capita income classes				
Number of regions	[0,75)	[75,90)	[90,110)	[110,125)	[125,∞)
46	0.78	0.09	0.13	0.00	0.00
45	0.33	0.43	0.24	0.00	0.00
46	0.07	0.17	0.59	0.15	0.02
24	0.00	0.04	0.21	0.71	0.04
36	0.00	0.00	0.14	0.17	0.69